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Document de Travail
Working Paper
2021-20

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Monetary autonomy of CESEE countries and nominal convergence in EMU: a cointegration analysis with structural breaks

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21st February, 2021

Abstract

This paper investigates the monetary autonomy of Central Eastern and South Eastern European countries with the Euro area. These countries are European Union Member States that have not adopted yet the Euro single currency. Despite high degree of convergence as measured by Maastricht criteria, four of them do no plan to enter the Euro area soon. We therefore assess monetary autonomy of these countries over the long run through the use of a multivariate cointegration methodology with structural breaks (Johansen et al., 2000). This methodology allows us to capture the multidimensional aspects of monetary autonomy in the context of nominal convergence in the Economic and Monetary Union, by including both domestic and Euro area variables into the system (policy rates, inflation rates, exchange rate). It also enables us to exploit all information contained in the macroeconomic series of these countries, for which broken economic history translates into non-stationary time series with breaks. Our empirical results suggest that modelling structural breaks changes the number and/or nature of cointegrating relations between our variables compared to the standard error correction model without breaks. With this modelling, we find monetary policy spillover from the Euro area to Bulgaria, the Czech Republic, Hungary and Romania. The inclusion of Euro area inflation to our baseline model enriches the cointegrating relations for the Czech Republic and Bulgaria. Poland is found to be the most monetary-independent country of our study across the various models estimated. On the other hand, Romania's monetary interdependence with Euro area is better modelled without taking into account any structural break.

JEL Classification: F31, F36, F42, P33
Keywords: CESEE, EMU, EU, nominal convergence, monetary autonomy, structural breaks, cointegration, integration

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Contents

1 Introduction

Central, Eastern and South Eastern European (CESEE) countries[1] share a common history as Socialist economies, as Transition economies after the end of the Soviet Union, more recently, as European Union (EU) New Member States and now as Euro area Accession countries. These six countries have in common several mechanisms of economy functioning, macroeconomic imbalances (particularly large public and private external debt) and financial vulnerabilities (exacerbated credit cycles, cross border banking relationships with the Euro area and sensitivity to the financial channel). Their development level measured by GDP per capita in purchasing power terms is around half the EU average level and the Gloabl Financial Crisis (GFC) has led to a slowdown of the rate of income convergence with the EU (Crespo-Cuaresma et al., 2012). Institutions standards are still in a catch-up process with OECD and EU countries (European Commission, 2018).

To be able to adopt the single currency, CESEE countries are required to fulfill some targets in terms of inflation, long-term interest rates, public debt, public deficit and nominal foreign exchange: these requirements are known as Maastricht Convergence Criteria, defined in 1992. Compliance with these requirements in terms of nominal and public finance convergence is monitored every two year by European Union institutions. As of June 2020, the European Central Bank (ECB) Convergence Report (European Central Bank, 2020) gives a mixed picture of nominal convergence of CESEE countries, because of a larger difference between fixed and floating-exchange rate countries since two years: regarding the price stability criterion, only fixed or quasi-fixed exchange rate Southern Eastern European countries (Croatia and Bulgaria) comply with the 12-month average inflation rate reference rate of 1.8%. On the contrary, Central European countries and Romania, that are in a floating or managed floating exchange rate regime, saw their inflation rates above or expected to rise above this reference level (annual inflation stands at 2.9%, 3.7%, 2.8% and 3.7% for the Czech Republic, Hungary, Poland and Romania respectively as of May 2020). With regard to the convergence of long-term interest rates, all countries recorded below the reference rate of 2.9%, except Romania (4.4%). As regards the fiscal criterion, four countries out of six are compliant with the deficit criteria; Hungary and Romania are under Significant Deviation and Excessive Deficit Procedures respectively. Croatia and Hungary are the only countries with an expected 2021 general government debt-to-GDP ratio exceeding the reference value of 60%. Despite an heterogeneous inflation situation, performance in terms of fiscal position and long-term interest rates remains fair.
If we finally consider the exchange rate criterion, that is the participation to the Exchange Rate Mechanism II (ERM II), it is noticeable that until July 2020 and despite compliance with Maastricht criteria in some cases, none of the CESEE countries participated. Bulgaria had to delay its application process to ERM II in June 2018 despite fulfillment of Maastricht criteria. The reason lies in the fact that accession criteria to ERM II were reinforced at this occasion by additional "prior commitments" in country-specific policy areas and with mandatory participation to the Banking Union's Single Supervision Mechanism (SSM) through "ECB close cooperation" and to the Single Resolution Mechanism -SRM- (Convergence Report 2020, box 1.4).

[1] Bulgaria, the Czech Republic, Croatia, Hungary, Poland, Romania.
ISO-3 country codes: BGR, CZE, HRV, HUN, POL, ROU

As of February 2021, Bulgaria and Croatia are now participant Member States to the ERM II since July 2020, after being applicant Member States since July 2018 and July 2019, respectively. On the contrary, the Czech Republic, Hungary, Poland and Romania do no plan to apply to ERM II in the near future. One may then wonder why CESEE countries delay their participation to ERM II, instead of choosing to deepen their already encouraging nominal convergence with the EMU as described above. Some answers may surely be found in the area of politics, with the rise of Euro scepticism and populism in all EU countries, with UK Brexit exacerbating such movements. Our research hypothesis is that CESEE countries have faced and continue to face macroeconomic policy issues and fear of losing two monetary policy instruments. Firstly, they fear of losing the policy rate to adjust to shocks, particularly to financial crises; secondly, the exchange rate interventions, both as a monetary policy instrument in a context of high exchange rate pass-through to inflation and as a financial stability instrument in the context of large currency mismatch in domestic balance sheets. This type of monetary regime combining inflation targeting and foreign exchange market interventions has been recently analyzed both empirically and theoretically in Ghosh et al. (2016).

The fear of losing monetary policy rate hypothesis is supported by the fact monetary policy responses have been extensively used by CESEE countries to adjust to financial shocks in the past decade. If we consider the Global Financial Crisis (GFC) in 2008, CESEE countries have been severely hit, with empirical literature documenting larger output losses in CESEE countries compared to other Emerging regions (Gallego et al., 2010; Allegret, 2012). National Central Banks (NCBs) have responded with monetary policy actions, although some of them have been delayed, procyclical or non-existing, pointing towards Fear of Floating phenomenon (Calvo and Reinhart, 2002): Hungary for instance increased its policy rate in October 2008 and Croatia did not adjust it. After this initial phase, CESEE countries used a large spectrum of unconventional monetary policies: if we refer to Ishi et al. (2009) taxonomy, all CESEE countries except Romania used a domestic mix of liquidity easing measures (reserve requirements, easing of collateral requirements) and Hungary, Poland and Romania used foreign exchange liquidity injections (Ishi et al., 2009; Magyar Nemzeti Bank, 2018). If we consider the Covid crisis in spring 2020, four NCBs started asset purchase programs which affect long-term yields (Magyar Nemzeti Bank, 2020) and the Hungarian Central Bank (MNB) has even responded with a revision of monetary policy instruments.
The fear of losing foreign exchange interventions as a monetary policy instrument is supported by the fact that NCBs both communicate around and use this policy instrument first to maintain price stability (pointing towards a high exchange rate pass-through to inflation, even though the effect has been decreasing as documented in the literature by Égert and Macdonald (2009)). For instance, the Czech National Bank used an exchange rate commitment from November 2013 until April 2017. NCBs also emphasize the interaction between monetary and macroprudential policies even though they theoretically have separate instruments (as stated in Czech National Bank's Inflation Report, for instance), suggesting foreign exchange interventions may be used as an instrument to smooth undesirable currency balance-sheet impacts. This is in line with Josifidis et al. (2014) conclusions that show, within a structural VAR framework, that currency mismatch in these countries explains the lags in monetary policy responses to the GFC and may lead to both fear of floating and fear of losing international reserves.

Last but not least, because they are EU Member States and part of the Economic Union with free capital movement, CESEE countries are also faced with the traditional monetary policy trilemma (Mundell, 1961), where two monetary policy goals can be achieved only between exchange rate stability, financial openness and monetary independence (defined as the capacity to manage short-term interest rates). They may also be constrained by the monetary dilemma (Rey, 2015) and the transmission effect of the global financial cycle that overcomes domestic monetary stance. CESEE countries' current degree of nominal convergence with the Euro area, combined with exchange rate stability and free capital movement point towards a lack of autonomy of monetary policy within the monetary policy trilemma and dilemma frameworks.

Given the motivations just developed, our research question is therefore to determine to which extend CESEE countries have monetary policy autonomy in the context of nominal convergence with the EMU. We test and estimate the monetary autonomy assumption of CESEE countries with the Euro area, making sure that our analysis matches the horizon over which convergence sustainability is assessed by European institutions. Hence our analysis focuses on long-term relations between domestic variables entering the reaction function of National Central Banks (inflation rates, industrial production, nominal exchanges rates with the Euro) and Euro area variables (ECB reference rate, Euro area inflation). To perform this analysis, we rely on individual country time-series analysis. We concentrate on Euro area conventional monetary policy impact on domestic monetary policy variables, based on the empirical evidence of monetary transmission mechanisms for these countries (see Égert and Macdonald (2009) for a literature review). Importantly, we do not focus on unconventional monetary policy spillovers because CESEE countries have not extensively used quantitative easing measures as in the Euro area since 2015 and literature on Euro area's unconventional monetary policy spillovers to CESEE countries is relatively scarce to understand their transmission channels.

One of the main contributions of our paper is to exploit all information contained in the macroeconomic series of CESEE countries, for which broken economic history translates into non-stationary time series. Data are characterized by breaks in levels and trends, and also contain stochastic trends. We therefore use cointegration methodology within a multivariate setup that specifically models structural breaks in the deterministic terms, both for the relations and the underlying time series. We use an extension of the Johansen rank test (Johansen et al., 2000) to determine the number of cointegrating relations. By using this methodology, we make the implicit assumption that the nature of the cointegrating relations between variables does not change over time[2], but that their levels and/or trends may change over time due to structural breaks. We therefore rely on standard macroeconomic models to interpret our results. Another contribution is that we assess the importance of the exchange rate in the conduct of monetary policy in these countries by including the nominal exchange rate versus Euro in our variables of interest. Our empirical findings are as follows: for all countries, structural break modelling has changed the number and/or nature of cointegrating relations between our variables compared to the standard VECM without breaks. For Bulgaria, the Czech Republic and Hungary, we find monetary policy dependence with the Euro area in a structural break model, in accordance with their monetary regime. Poland is quite independent from the Euro area, with results robust across various models. On the contrary, the structural break model is not suited for Romania and Croatia, even though we find Euro area monetary spillover to Romania in a standard VECM model. The inclusion of Euro area inflation to our baseline model enriches the cointegrating relations for the Czech

[2]There is no notion of "regime" such as in the low-volatility/high volatility regimes

Republic and Bulgaria. We also find that the exchange rate variation is not long-term restricted for most cointegrating relations, which potentially validates the importance of the exchange rate in the conduct of monetary policy. The paper is organized as follows. Section 2 highlights some facts and reviews the empirical literature about the monetary autonomy of CESEE countries with the Euro area. Section 3 presents the challenges inherent to the data and the subsequent modelling approach. Section 4 gives an overview of the cointegration methodology in the presence of structural breaks. Section 5 focuses on the empirical framework. Section 6 comments the main results and Section 7 concludes.

2 Monetary autonomy of CESEE countries with the Euro: facts highlight and literature review

We first focus on the long-term dynamics of domestic monetary policy rates. Then, we review the EU institutional nominal convergence framework within which CESEE countries formulate their domestic monetary policies, which implies these countries have domestic targets and EMU convergence requirements for customer price inflation and nominal exchange rates (with econometric properties of these series formally analyzed in Section 3). Finally, literature review highlights the main strands of empirical research to which our paper is related.

2.1 Monetary policy rates dynamics

Since 1991, CESEE countries have experienced major macroeconomic changes, switching from socialist to small open economies in less than 20 years. These countries now account for a large range of exchange rate regimes (from currency board arrangement to free floating exchange rates). Regarding the type of nominal anchor, CESEE countries are split between countries currently following an inflation-targeting monetary policy framework (the Czech Republic, Hungary, Poland and Romania) and countries using the exchange rate as the nominal anchor (Bulgaria and Croatia).
If we analyze the patterns of domestic and Euro Area central bank policy rates over the long run, we can distinguish between two periods: one divergent and one convergent, with speed of convergence depending on the country considered. For inflation-targeting countries (Figure 1), domestic policy rates share a rapidly decreasing trend during the transition period, characterized by liberalization of the economies and a strong disinflation, up to 2001-2002 for the Czech Republic, Poland and Hungary and up to 2005 for Romania. It is noticeable that for these countries, inflation-targeting regimes have been put in place despite high inflation levels. This period is followed by a flattening/stabilization period when both Euro area and CESEE policy rates seem to follow a similar pattern.

For exchange rate-targeting countries (Figure 2), we have two different cases. Bulgarian policy rate closely mirrors the dynamics of Euro area policy rate (which is explained by the currency board fixed exchange regime, which makes the policy rate endogenous to monetary policy); it is not the case for Croatia policy rate, which has common trends but much higher levels compared with the Euro area policy rate.

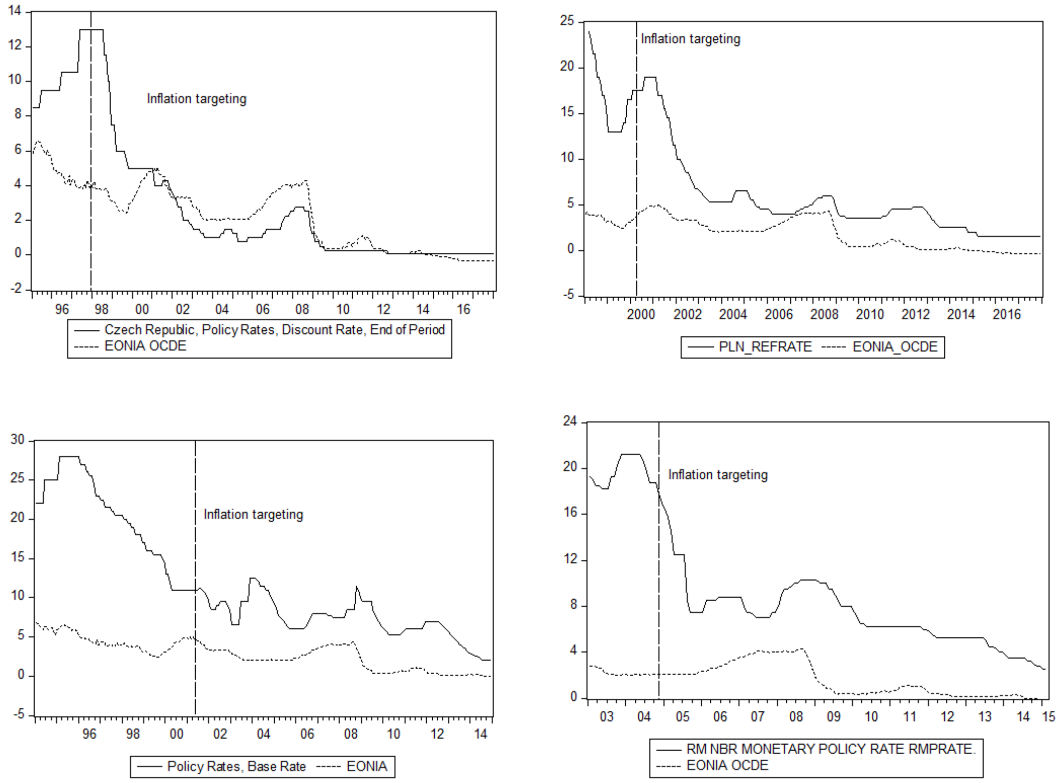


Figure 1: Inflation-targeting countries - Policy rates evolution
Top left/right: CZE/POL ; Bottom left/right: HUN/ROU

Notes: Expressed in percent; "Inflation-targeting" dashed line indicating starting year of IT monetary regime; Data source: OECD and National Central Banks

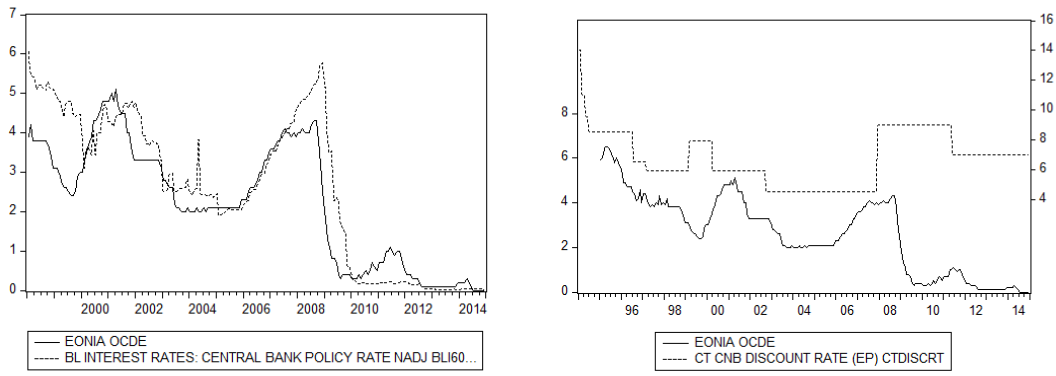


Figure 2: Exchange rate-targeting countries - Policy rates evolution
Left figure: Bulgaria ; Right figure: Croatia

Expressed in percent points; Data source: National Central Banks

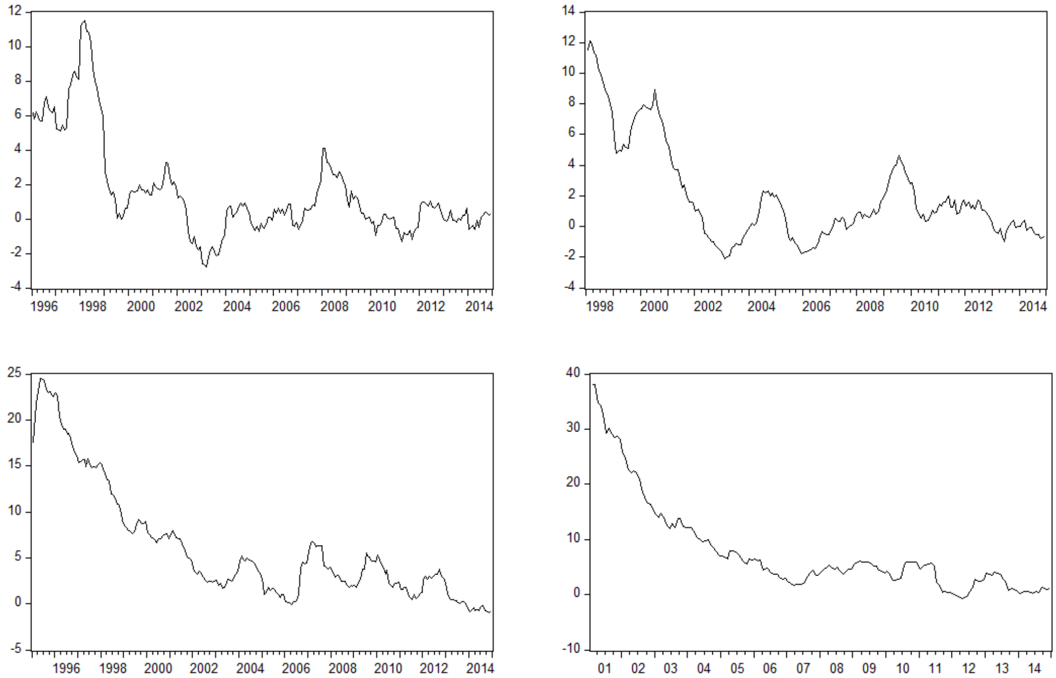


Figure 3: Inflation-targeting countries - Inflation rate differential with Euro area
Top left/right: CZE/POL ; Down left/right: HUN/ROU

Notes: Expressed in percent points; Data source: National Central Banks

2.2 Domestic policy implications of EU nominal convergence framework

CESEE National Central Banks have a dual requirement to achieve monetary convergence with the EU: price and exchange rate stability as defined by the Maastricht Treaty. The price stability criterion states the rate of consumer prices inflation must not be more than 1.5% higher than the average of the three best performing EU countries. The exchange rate stability criterion is fulfilled if the exchange rate moves within a +/-15% fluctuation band, without devaluation of the central rate and excessive pressures on the exchange rate.

Even though these criteria are not translated into official monetary policy objectives by National Central Banks, they still have potential implications in terms of domestic monetary policy. As regards price stability, CESEE countries have an implicit dual inflation target, i.e. one in terms of national CPI and one in terms of Harmonized CPI. This institutional situation may impact the inflation target set by monetary authorities and lead to enhanced Euro area inflation transmission to domestic inflation. Exchange rate interventions may also be used as an additional instrument to maintain price stability, in case the exchange rate pass-through to inflation is strong (this framework being a "flexible Inflation Targeting" as in the Czech Republic). The Czech Republic has for instance put in place an exchange rate floor of 27 korunas per euro in November 2013 to stop koruna appreciation and help push inflation towards its 2% target. As far as exchange rate stability is concerned, it may require a second monetary instrument to be achieved, for instance interventions on the foreign exchange markets (following the rule under which there should be one instrument per policy objective). This is actually the case for the Czech, Hungarian and Romanian central banks that allow sterilized foreign exchange market interventions as a monetary policy instrument (European Commission, 2018).

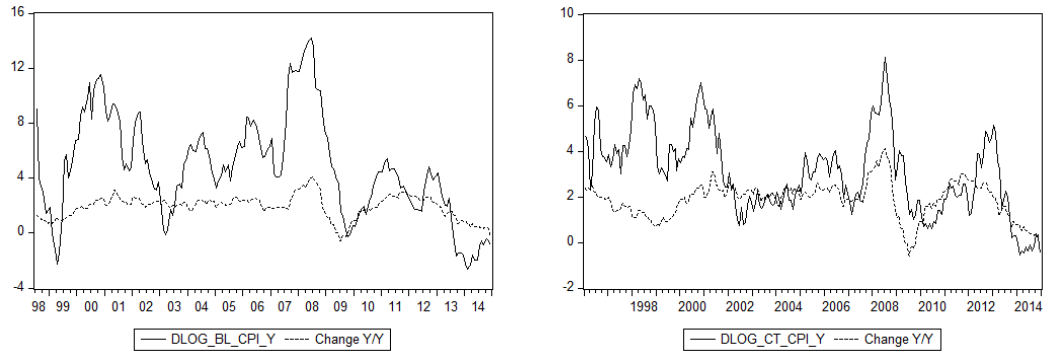


Figure 4: Exchange rate-targeting countries - Inflation rate evolution
Left figure: Bulgaria / Right figure: Croatia

Notes: Expressed in percent points; Dashed line: Euro area inflation rate; Data source: National Central Banks

2.3 Price and exchange rate dynamics

Long-run graphical analysis of inflation dynamics between Euro area and domestic countries in Figure 3, measured by the domestic minus Euro area CPI inflation differential and in Figure 2, shows that there is a large synchronization of inflation rates starting in the 2000s for the Czech Republic, Poland, Croatia and post GFC for Hungary and Romania. One noticeable exception is Bulgarian inflation, which has fluctuated within a large range and does not show a large degree of convergence with Euro area inflation. This convergence process has been precedented by a long disinflation period, starting in 1991 and lasting for a decade for most countries under review. In a similar fashion as for inflation, CESEE nominal exchange rates[3] have been converging towards more stable values compared to Euro over the past decades (Figures 5 to Figure 10). We can separate our countries of analysis depending on their current foreign exchange regime. If we look at "floaters" countries first, domestic exchanges rates have followed a common pattern of managed nominal depreciation until the beginning of the 2000s (in most cases, through crawling peg arrangements). Czech koruna has appreciated since, coinciding with a managed floating system, with its appreciation stopped with the ceiling put in place in 2013. Polish zloty, Romanian leu and Hungarian forint have globally depreciated after free floating was instaured (in April 2000, November 2004 and February 2008, respectively). If we now look at "fixers" or "tightly managed exchange rate" countries, Bulgarian lev exchange rate is fixed to the Euro within the framework of a currency board since June 1997. Croatian kuna is pegged to the Euro through a tightly managed floating exchange regime (crawling band) since the beginning of the 2000s.

[3]Expressed as one unit of domestic currency per Euro

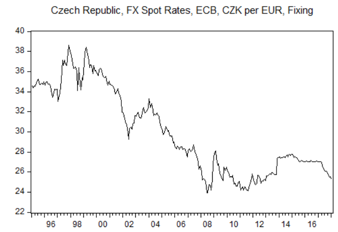


Figure 5: Czech Krona

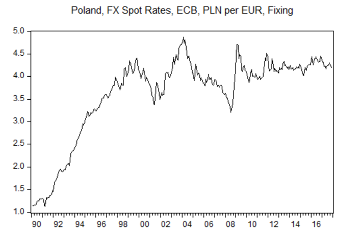


Figure 6: Polish Zloty

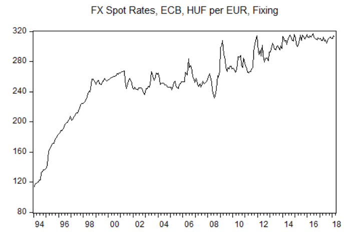


Figure 7: Hungarian Forint

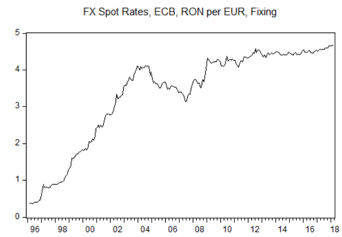


Figure 8: Romanian Leu

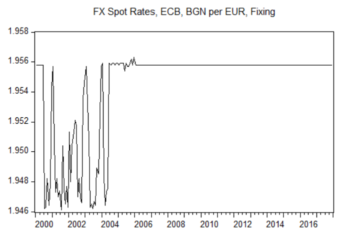


Figure 9: Bulgarian Lev

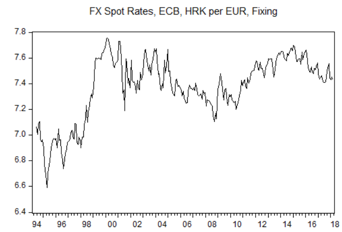


Figure 10: Croatian Kuna

2.4 Monetary autonomy of CESEE countries in the empirical literature

Our paper primarily estimates the monetary independence of CESEE countries with the Euro area and is as such related to ECB monetary policy spillover literature on CESEE countries. This literature is not unified because it uses a large spectrum of specifications and econometric methodologies, with metrics that can be quite different. In line with the larger EMU Accession countries monetary integration literature (Camarero et al., 2002; Fountas and Wu, 1998; Frömmel and Kruse, 2015), the question of monetary policy autonomy of CESEE countries has been commonly assessed using an univariate monetary spillover equation (derived from Uncovered Interest Parity -UIP-)[4], with cointegration being the measure for monetary integration (see amongst others Brada et al. (2005); Holtemöller (2005); Kasman et al. (2008)). One shortcoming as highlighted in Camarero et al. (2002) is that relying on UIP to test for convergence in the EU means to assume that forecast errors, risk premium and exchange rate changes are all stationary processes. Instead of interest rates transmission, several papers analyse volatility transmission of interest and exchange rates (Crespo-Cuaresma and Wójcik, 2006; Habib, 2002; Windberger et al., 2012). A more recent literature analyses Euro area unconventional monetary policy spillover on CESEE countries (Bluwstein and Canova, 2016; Hajek and Horvath, 2018; Moder, 2017).

There is therefore a gap in the literature assessing CESEE countries' monetary policy autonomy in a multivariate framework, which has been predominant in the domestic monetary transmission literature, but within a closed economy setting (see Égert and Macdonald (2009) for a literature review). Many papers were written in the beginning of the 2000s and analyzed new EU members, namely Central Europe countries (Poland, Czech Republic and Hungary) together with Baltic countries, Slovakia and Slovenia. Standard econometric methodology is a Vectorial Auto Regressive (VAR) representation and estimation, with either a recursive or structural identification of shocks (Creel

[4] Assuming the equivalence between monetary spillovers and lack of monetary autonomy

and Levasseur, 2006; Elbourne and de Haan, 2009; Matei and Héricourt, 2006). In the perspective of EMU though, some articles have been modelling the European Monetary System (EMS) context and evaluating cross-country differences in monetary policy transmission mechanism within the VAR framework. In Peersman and Mojon (2001), authors form three groups of countries (from 10 Euro Area core members) depending on their monetary regime within the EMS. They include both German policy rate and Deutsche Mark bilateral exchange rate in the VAR and define for each group an ad hoc identification scheme for the monetary policy shock. They find real GDP and inflation responses to German monetary policy shock that are consistent with expected effects. We try to fill the gap in the literature by extending the Peersman and Mojon (2001) approach to Central, Eastern and South Eastern European countries and assessing monetary autonomy of CESEE countries within a multivariate cointegrated VAR system. We embed the relevant foreign variables (exchange rate and Euro Area inflation) and quite importantly, model structural breaks which allows us to tackle the challenges in CESEE macroeconomic series (developed in subsequent Section 3). Modelling structural breaks should be a control for the price puzzle[5]. To our knowledge, there is only one article using cointegration techniques with structural breaks like we do in this paper in a multivariate setting: Koukouritakis et al. (2014) that investigate the monetary transmission mechanisms through interest rate and real effective exchange rate channels for five South-Eastern European countries. Their analysis is performed in a closed economy framework and they do not test the long-run zero restrictions.

Given the wide spectrum of exchange rate regimes of CESEE countries, our paper is also related to the literature on the role of exchange rates regimes in the international transmission of interest rates. Frankel et al. (2004); Obstfeld et al. (2004) have recently strongly contributed to this debate for Emerging countries. Frankel et al. (2004) cannot reject the full transmission hypothesis of interest rates even for open economies with a flexible exchange rate. It therefore suggests countries do not have as much monetary autonomy as theory implies. Căpraru and Ihnatov (2012) amongst others also contributed to this debate for CESEE countries, while some additional empirical literature appeared in the aftermath of the GFC with an analysis of international real, financial and monetary shock transmissions (Allegret, 2012; Jiménez-Rodríguez et al., 2010; Josifidis et al., 2014). Main takeaways are that domestic monetary policies following the GFC have been determined by the exchange rate regimes (fixed or floating), and more generally by the Fear of Floating (Calvo and Reinhart, 2002). Based on these empirical findings and the current exchange rate policies adopted by National Central Banks (NCBs) as exposed in Section 2, we are interested in assessing the importance of the exchange rate in the reaction function of the NCBs of CESEE countries. We therefore embed nominal exchange rate as one of the endogenous variables of our model (Aizenman et al., 2011; Ball, 2000; Taylor, 2001; Wollmershäuser, 2006). To our knowledge, the inclusion of exchange rates in the empirical CESEE monetary literature has been followed by Frömmel et al. (2011) who estimated monetary policy rules within an augmented Taylor rule setting, with same methodology as in Gerlach-Kristen (2003).

[5]The price puzzle (an increase of prices following an increase in the interest rate) is often reported in CESEE monetary VARs and has been extensively analyzed in the literature. Peersman and Smets (2001) get expected responses from output and prices following a monetary shock by adding the foreign exchange and the German interest rate. Égert and Macdonald (2009) highlight that modelling breaks helps solving the price puzzle.

3 Challenges in the data and modelling approach

3.1 Broken macroeconomic history translates into broken macroeconomic series

As mentioned previously, CESEE countries have switched from socialist to small open economies in less than 20 years. Exchange rate regimes have evolved, starting from fixed exchange rates to a various range of exchange rate regimes (floating, soft pegs, hard pegs). Monetary regimes have also changed a lot, moving from exchange rate targeting to inflation targeting regimes. Table 1 summarizes the main structural changes in terms of monetary and exchange rate regimes, together with the EU membership dates. All these dates are exogenous break dates as they are officially announced by national authorities. It is noticeable that CESEE countries have gradually changed their monetary and foreign exchange regimes over a rather short length of time. As commented in Section 2, this broken domestic macroeconomic history translates into macroeconomic series that are characterized by broken trend and levels. Global or symmetric shocks such as the GFC also have an impact on domestic macroeconomic series.

In terms of methodology, these structural breaks matter for CESEE countries for two main reasons:
- Structural breaks change critical values of unit root and cointegration tests. Since our paper is focused on long-term relations between monetary variables, structural breaks are therefore to be taken into account carefully.
- Compared to advanced countries for which there exists long data series, CESEE countries have relatively short time series (starting end of the 1980s). Accounting for structural breaks (in the deterministic terms) allows to avoid the "split sample" estimation solution that is suited for large finite samples.

Table 1: Main structural break dates: monetary regime, exchange rate regime, European Union membership

Country	Monetary and exchange rate regimes	Start date	End date
Bulgaria	Exchange rate anchor — Hard peg (currency board)	**1997:M7**	...
	EU membership	**2007:M1**	...
Croatia	Exchange rate anchor — Soft peg (stabilized arrangement*)	**1994**	...
	EU membership	**2013:M7**	...
Czech Republic	Exchange rate anchor — Conventional fixed exchange rate	1993:M1	1996:M2
	Transitional inflation targeting — Soft peg (corridor)	1996:M2	1997:M12
	Inflation Targeting — Managed floating	**1997:M12**	2013:M11
	EU membership	**2004:M5**	...
	Inflation Targeting — Floating with exchange rate commitment	2013:M11	2017:M4
	Inflation Targeting — Floating with foreign exchange interventions	2017:M15	...
Hungary	Exchange rate anchor (Crawling peg)	1995:M5	2001:M5
	Inflation Targeting — Crawling band	**2001:M5**	2008:M2
	EU membership	**2004:M5**	...
	Inflation Targeting — Floating with with foreign exchange interventions	2008:M2	...
Poland	Transitional inflation targeting — Crawling peg	1995:M5	2000:M4
	Inflation Targeting — Free floating	**2000:M4**	...
	EU membership	**2004:M5**	...
Romania	De jure Managed floating regime	1991	2004:M11
	Inflation Targeting — Stabilized arrangement*	**2004:M11**	...
	EU membership	**2007:M1**	...

Source: Habib (2002); Josifidis et al. (2014), National Central Banks. "..." indicate that regime is still current.
Definitions and latest exchange rate arrangement indicated as per IMF's Annual Report on Exchange Arrangements and Exchange Restrictions (2019)
*Exchange rate crawling band **The country maintains a de facto exchange rate anchor to the euro

3.2 Accounting for structural breaks when modelling CESEE macroeconomic series

Structural breaks have been a challenge for academic research because they complexify the determination of stochastic or deterministic trends in macroeconomic time series. One important result from the unit root econometric research has been to establish that standard unit root tests' ability to reject the unit root null hypothesis in the presence of structural breaks decreases. In his seminal paper, Perron (1989) challenged the empirical results of the Nelson and Plosser (1982) analysis of American time series over a long period, in which they showed that macroeconomic series had a unit root. To support its analysis, he uses a modified Dickey-Fuller unit root test that includes dummy variables to allow for one known break. Perron (1989) test has been extended in many ways, first by using two types of models (Innovative Outlier versus Additive Outlier models, see Perron and Vogelsang (1992)), secondly by using an endogenous procedure to estimate break dates (Zivot and Andrews, 1992), thirdly by allowing several structural breaks in the data (Lumsdaine and Papell, 1997) and finally, by accounting for the existence of structural breaks both under the null and the alternative hypotheses (Clemente et al., 1998; Lee and Strazicich, 2003). Most tests model breaks under the alternative hypothesis and not under the null (or not the same number of breaks under H0 and H1).

In order to determine the degree of integration of CESEE countries' variables, we therefore use the Clemente et al. (1998) unit root test that tests the null hypothesis of a unit root with possibly one endogenous break in mean, against the alternative of no unit root with two endogenous breaks[6]. This test is an extension of the Perron and Vogelsang (1992) unit root test. It is the first test that allows to detect non-stationary variables with breaks under the null and also allows for the existence of two structural breaks. The test does not model the deterministic trend, which is in line with Lütkepohl (2004) that indicates it may be desirable to subtract the deterministic term first when the stochastic part is of primarily interest for econometric analysis. Moreover, this test is parametric so it is important to correctly determine the lag truncation parameter, amongst other parameters.

Table 2 summarizes unit root test results on main domestic macroeconomic series of the six CESEE countries (namely, industrial production, domestic inflation rates, policy rate, exchange rate annual variation). We do not reject the null hypothesis of all series having a unit root and at least one endogenous break date at 5% level.

For Central European countries (Poland, the Czech Republic and Hungary), first break date for all series is located at the end of the 1990s (between 1999 and 2001) and corresponds to the implementation of inflation-targeting monetary policies. Second break date is either related to the GFC (2008 up to 2010), or between 2001 and 2004, which corresponds to the years preceding their EU membership in April 2004. For Southern Eastern European countries (Bulgaria, Romania and Croatia), endogenous break dates correspond to either the start of the inflation-targeting regime (eg Romania), the EU membership accession (eg Bulgaria for the nominal variables), or the years following the GFC (for all three countries, between 2009 and 2011). Croatia also experienced a break in its domestic policy rate and its exchange rate variation during the course of the year 2000. Finally, for the Euro area, both policy rate and inflation rates exhibit break dates corresponding to the impact of the GFC (2009-2010), and another interesting break date either in 2001 or in 2006. This reflects the fact that the start date of the analysis determines which breaks are captured: since we want to determine

[6] See Appendix A for test hypotheses and statistics

Table 2: Clemente et al. (1998) unit root test with double mean shifts, Additive Outlier model

Country	Variable	Lag truncation	Optimal break point 1	Optimal break point 2	min T-Stat
Poland	IPI	1	2003:M9	2009:M12	-3.71
1998:M3	Δ_CPI	12	2001:M12	2008:M3	-2.18
	i	18	1999:M11	2001:M3	-1.15
	Δ_e	12	2009:M3	2009:M7	-2.88
Czech Republic	IPI	24	2003:M5	2014:M6	-3.21
1996:M1	Δ_CPI	2	2001:M11	2008:M12	-5.00
	i	2	1998:M9	2002:M4	-3.14
	Δ_e	2	2007:M1	2010:M1	-2.51
Hungary	IPI	22	1998:M12	2003:M6	-2.63
1994:M1	Δ_CPI	24	1997:M6	2001:M12	0.03
	i	0	1999:M3	2005:M4	-2.89
	Δ_e	24	1996:M4	1999:M6	-2.60
Romania	IPI	12	2006:M6	2012:M4	-2.77
2003:M1	Δ_CPI	12	2005:M11	2011:M10	-3.88
	i	17	2005:M5	2010:M4	-0.94
	Δ_e	12	2008:M2	2009:M8	-2.48
Bulgaria	IPI	5	2004:M12	2008:M7	-2.54
2001:M1	Δ_CPI	21	2007:M10	2009:M4	-1.66
	i	23	2007:M4	2009:M2	-1.85
	Δ_e	24	2004:M3	2005:M7	-4.39
Croatia	IPI	4	2003M11	2010:M7	-2.52
1997:M6	Δ_CPI	24	2008:M1	2009:M1	-1.48
	i	13	2000:M6	2010:M2	-3.63
	Δ_e	23	2000:M3	2008M9	-4.48
Euro area					
1994:M1	i*	18	1996:M11	2009:M5	-1.45
1996:M1	i*	11	2001:M2	2009:M5	-2.28
1998:M3	i*	13	2001:M2	2009:M8	-2.15
2001:M1	i*	5	2006:M11	2009:M3	-3.07
2003:M1	i*	1	2006:M9	2009:M3	-6.82
2001:M1	Δ_CPI	16	2009:M2	2010:M4	-1.27
2003:M1	Δ_CPI	16	2009:M2	2010:M4	-1.18

Notes: Critical value @ 5% = -5.49 ; critical value for k(t) AO: Additional Outlier. The k(t) method for determining the lag truncation parameter is the one used in Perron and Vogelsang (1992). Trimming parameter lambda has been chosen to 5% which is line with article recommendations to have largest data window possible.

global breaks more than local breaks, we will use the longest data spam available. We perform a robustness analysis of this unit root test as regards lag truncation determination and innovation models used for the endogenous determination of break dates in Subsection 6.4. Endogenous break dates are robust to the innovation model used because of the quite large number of lags used.

3.3 Introducing multivariate cointegration methodology with structural breaks for CESEE countries

In terms of methodology, two main types of cointegration models have been initially considered to answer our research questions.

i) Multivariate cointegration with structural breaks has many economic and methodological advantages. First of all, CESEE monetary convergence process is multidimensional (inflation, foreign exchange, short-term interest rates). It is therefore best modelled by a multivariate cointegrating system. As highlighted by Perron (1989), cointegration relations with structural breaks in the deterministic terms are good model candidates for macroeconomic analysis. This conclusion, applicable for advanced countries with stable institutions and macroeconomic regimes, may be questionable for Emerging countries; but we believe it applies to Emerging Europe since it has been engaged in a convergence process with the European Union following their independence at the beginning of the 1990s. Moreover, since we assume no structural change in the long-run relations, we can therefore use international relationships to analyze our results, such as the Uncovered Interest Parity (UIP) framework.

ii) Threshold cointegration (either in the short-term dynamics or in the cointegrating vector) has been a model candidate but was finally put aside for our countries for the following reasons. First of all, cointegration tests do not allow to test series displaying deterministic breaks (cf the test regressions for the Gregory-Hansen cointegration test in Gregory and Hansen (1996)). Second reason is that seminal estimation model for threshold cointegration (Hansen and Seo, 2002) is a bivariate one. We have estimated one relationship between domestic and ECB policy rates that was very difficult to interpret, even though the model did fit the dynamics of the variables (results available on request). Finally, multivariate threshold models use exogenous thresholds (and not the error correction term). Such choice of exogenous threshold variable is not easy between domestic and international variables.

Section 4 gives theoretical background on the model used in this paper, that is a cointegration model with breaks in the deterministic terms as developed by Johansen et al. (2000). Model is suited for up to two exogenous breaks, which are breaks that are not estimated. Main difference of this model with the one developed by Johansen (1995) relates to asymptotic null distribution of the rank test statistic, because the introduction of shift or trend dummies modifies critical values.

4 Cointegration analysis in the presence of structural breaks

4.1 Representation of the VECM with structural breaks in trend and level

General framework for cointegration analysis is the one from Johansen et al. (2000). The error correction formulation of the vector autoregressive model is:

$$\Delta X_t = \alpha \left(\beta' X_{t-1} + \gamma' t\right) + \mu + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \epsilon_t \tag{1}$$

Where X_t is the vector of p variables at time t, ΔX_t is the same vector in first differences, α and β full rank matrices of dimension $p \times r$ (α being the long-term adjustment matrix, β being the cointegrating matrix), r the rank of the matrix $\alpha\beta$', γ the p-vector of deterministic trend (drift) parameters, μ the p-vector of constant (mean) parameters, k the number of lags, ϵ the p-vector of independent and identically distributed errors (iid) with mean 0 and variance ω (Gaussian white noise vector).

This initial framework is extended to account for structural breaks in deterministic terms (mean and trend). The original article presents three models, but the most suited to our data and research objective is the $H_l(r)$ model, following article's notations. The data-generating process has the following trend-restricted VECM representation:

$$\Delta X_t = \alpha \binom{\beta}{\gamma}' \begin{pmatrix} X_{t-1} \\ tE_t \end{pmatrix} + \mu E_t + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \sum_{i=1}^{k} \sum_{j=2}^{q} k_{j,i} D_{j,t-i} + \epsilon_t \tag{2}$$

With the same notations as in Equation 1 except for q the number of samples, $E_t = (E_{1,t}, ..., E_{q,t})$ the q-vector of sample dummies, γ the $q \times r$ matrix of deterministic trend (drift) parameters, μ the $p \times q$ matrix of mean parameters (μ_j varying across samples, from 1 to q), $D_{j,t-i}$ the indicator function for the ith observation in the jth sample (corresponding to X_{Tj-1+i}), $k_{j,i}$ the p-vectors of dummy parameters (k dummies for each sample to condition the likelihood function over these initial observations at each subperiod).

Model $H_l(r)$ generates a multidimensional process X_t with a linear trend since ΔX_t has an unrestricted constant term $\mu \mathrm{E}_t$ (which differs between samples). Moreover, the cointegrating relations, $\beta' \mathrm{X}_t$, are trend stationary. Trend is restricted to the cointegrating vector (by means of the term tE_t, which also differs between samples) so that time series in X_t do not exhibit quadratic trends but only a drift. Hence, model $H_l(r)$ allows both non-stationary series and cointegrating relations to exhibit broken linear trends (both in mean and drift). This model is therefore suited for our series that potentially present structural breaks.
A process satisfying the rank hypothesis of $H_l(r)$ can be interpreted using Granger representation theorem: linear combinations of the process X_t, given by β, cointegrate while the process exhibits a linear trend in each of the subsamples. In vectorial terms, it means that the process $\beta' \mathrm{X}_t + \gamma' \mathrm{tE}_t$ is stationary for each sample period and has no trending behavior.

4.2 Johansen test for cointegration rank (Johansen et al., 2000)

The cointegration rank can be tested by modifying the procedures suggested by Johansen (1995). The statistical analysis is hardly changed but new asymptotic distributions arise, described by response surface analysis. To construct the test statistic, the likelihood function is maximized using canonical correlation methods applied to cointegration analysis by Johansen (1995). Inference is based on squared sample canonical correlations $\hat{\lambda}_i$ of ΔX_t and (X'_{t-1}, tE'_t)', corrected for ΔX_{t-i}, $D_{j,t-i}$ and tE_t.
The Likelihood Ratio test statistic for the $H_l(r)$ null hypothesis of at most r cointegrating relations against a general alternative $H_l(p)$ has the following form:

$$\mathrm{LR}\{\mathrm{H}_l(r)/H_l(p)\} = -T \textstyle\sum_{i=r+1}^{p} \log(1 - \hat{\lambda}_i) \tag{3}$$

In order to estimate the rank of $\alpha\beta$', a sequential testing procedure is necessary. The asymptotic distribution of the test only depends on the relative length of the sample periods, not on their ordering. Since no analytical solution is derived, first two moments of the Gamma distribution are approximated by simulation.
After the rank is determined, the model allows to test further restrictions on the parameters, particularly hypotheses on the slope of the deterministic trends. Authors find that tests for linear restrictions on both the slope for the cointegrating relations and the slope for the entire process are asymptotically χ^2 distributed. These two tests should

be performed sequentially. These tests are of particular interest for us since we can test whether our modelling with breaks in deterministic trends is statistically significant and serves our cointegration analysis.
Regarding model estimation, since there are no short-term constraint nor exogenous variables, it is done with Johansen (1995) Reduced Rank Maximum Likelihood procedure. Estimates of β are the r linear combinations of the data which have the largest empirical correlations with the stationary process ΔX_t. Even though modified Johansen test is very sensitive to the number of lags included in the test specification (due to its parametric nature), it is clearly the best suited to our macroeconomic data in terms of deterministic trend modelling and we therefore used it to determine the rank of our systems. Appendix B summarizes the alternative rank tests and the robustness checks of the modified Johansen rank test results.

5 Empirical methodology

5.1 Baseline model

Our multidimensional X_t process analyzed with the VECM with breaks consists of five variables of interest and the restricted trend term:

$$X_t = \{logIP_t, \Delta CPI_t, i_t^*, \Delta e_t, i_t, tE_t\}' \tag{4}$$

With $logIP_t$ the domestic industrial production transformed in logarithm, ΔCPI_t the annual customer prices inflation, i_t^* the Euro area policy rate, Δe_t the annual exchange rate variation against Euro, i the domestic policy rate and tE_t the deterministic term for each subsample.

We basically augment the closed economy model used in Peersman and Smets (2001) by adding Euro area policy rate as an endogenous variable. It is ordered third in the system because we do not want to normalize our cointegrating vectors on the variable we want to test. By doing so, we do not impose exogeneity to this variable and allow to test for the foreign policy endogeneity in the system.
This model is the one used in Peersman and Mojon (2001) for a VAR analysis that we follow in terms of variable ordering: domestic policy rate is ordered last, which means it is contemporaneously impacted by all other macroeconomic variables (output, inflation, Euro area policy rate and exchange rate). With the exchange rate ordered before the domestic policy rate, we assume the exchange rate has a contemporaneous impact on domestic monetary policy rate. With the Euro area policy rate ordered before exchange rate and policy rate, we assume Euro area monetary policy is not affected contemporaneously by domestic monetary policy nor exchange rate (which is in line with small open economy theory). Finally, domestic output and inflation are contemporaneously exogenous both to the Euro policy rate and to the domestic policy rate. We finally follow Peersman and Smets (2001) and do not include a money aggregate, because we assume the exchange rate monetary objective is stronger than the money aggregate one.
Regarding deterministic dummies, our specification includes: three sample dummies E_t^i, for i=1 to 3, representing the periods between the two breaks, that are used to model breaks in levels and also linear trend breaks when they are multiplied by the deterministic trend term tE_t^i (so, 6 deterministic terms in total).
Also, to make transition happen between the subperiods, Johansen et al. (2000) suggest

conditioning the log-likelihood function upon the first k observations of each sample period (which has the effect of setting the first k residuals to zero), with the transition periods being extendable if necessary. k dummies are therefore included at the beginning of each sample period.
All these deterministic dummies have to be included in order to avoid model misspecification, and particularly the transition dummies that capture the structural break outlier effect on the cointegrating relations.

5.2 Data overview

For national data, we use data from National Central Banks (policy rates), European Central Bank (exchange rates), National Statistical Offices and IMF IFS databases (for consumer prices and industrial production, respectively). Data are monthly. They are seasonally adjusted and transformed into natural logarithm except for interest rates. Data variation is calculated as year-on-year change. More precisely, domestic inflation rates are calculated as 12-month log differences between national Consumer Prices Indices, all items, after seasonal adjustment. Real activity is proxied by Industrial Production Indices (all items), that are taken in levels, seasonally adjusted and transformed into their natural logarithms. Exchanges rates are based on bilateral nominal exchange rate against Euro. Data are expressed in percentage points, except industrial production.
We pay special attention to the choice of the domestic policy rate, since there have been changes and/or addition of monetary policy instruments over the whole period (for instance, the introduction of Central Bank deposit rates or the Open Market Operations reference rates). We therefore use the main monetary policy instrument over the most recent period. For inflation-targeting countries (Hungary, Czech Republic, Poland and Romania), we use main monetary policy rate, which may differ across countries. They can be yields on money market bills, repurchase agreement rates, discount rates or deposit rates. They reflect the development stage of the financial markets and monetary institutional framework of these economies. For exchange rate-targeting countries (Bulgaria, Croatia), we use base rate for Bulgaria and the 3-month Zibor rate for Croatia as a proxy for the policy rate. We do not include the exchange rate as variable for Bulgaria because of the fixed nature of the exchange rate regime (and the stationarity of the series). We include it for Croatia, even though nominal anchor of monetary policy is the stability of the Kuna exchange rate against the Euro.
For Euro area, we use Euro Over Night Index Average (EONIA) rate (19 countries) from the OECD database, for which data is available since 1994[7]. This choice is motivated by the data spam that is much longer and is a good proxy for the ECB main policy rate[8].
Regarding the beginning of the analysis, we use longest spam of data available for each country. We end our analysis in December 2014 as common end date for all countries, so that the ECB Quantitative Easing period is not embedded into the data (it has officially started in 2015:M3).

[7]From January 1999, EONIA was computed as a weighted average of overnight unsecured lending transactions in the European Union and the European Free Trade Area (EFTA) interbank market prior to 1st October 2019.

[8]Refinancing rate is available since January 1999. We have compared EONIA with the various Euro area policy rates in Annex H.

Table 3: Start date of analysis - by country

Country	Analysis start date	Shortest data series and description of policy rates
Poland	March 1998	Reference Rate (minimum intervention rate on money markets and main instrument of monetary policy)
Czech Republic	January 1996	Customer Price Index Main policy rate (currently the 2-week repo rate)
Hungary	January 1994	Industrial Production (IMF) Central Bank Base rate series starts in 1990 (currently the 3-month deposit rate)
Romania	January 2003	Industrial Production (IMF) Monetary policy rate (currently the 1-week repo rate)
Croatia	June 1997	3-Month Zibor Repo rate series (current policy rate) starts in April 2005 only (Discount rate starts in 1992 but is not a policy instrument anymore)
Bulgaria	January 2001	Industrial Production national series changed in 2000

5.3 VECM specification

As exposed in Section 4, deterministic terms are included so that they correctly model the time series dynamics: we follow model $H_l(r)$ in which both non-stationary times series and the cointegrating relations can include breaks in the linear trend term. For break dates, we use the endogenous break dates as determined by the two-break unit root test and then proceed with the Johansen et al. (2000) cointegration test. Chosen break dates are either the policy rate break dates, either the breaks date that are common to most variables and are unchanged across models to allow comparability. This approach is standard practice even though using endogenous break dates in such a setting is discussed by Trenkler et al. (2008). It may be unclear whether the type of breaks (either pure deterministic or change in dynamics) is sufficiently known to make the tests applicable.

Since Johansen rank test is based on a Maximum Likelihood Gaussian function, it assumes residuals follow a multivariate normal distribution. We therefore perform Jarque-Bera normality test on residuals. Johansen rank test is parametric and is as such sensitive to the number of lags embedded to control for error autocorrelation. Hence, we have to determine the relevant number of lags to whiten the residuals[9]. The maximum lag analysis is performed through estimation of the unrestricted VAR model (inclusive of a constant and a trend). We use two types of approaches to select the lag order: the minimal information criteria (particularly the Hannan-Quinn criteria if there is no consensus) and the computation of several autocorrelation tests on the residuals for the multivariate system (Portmanteau test, Beutsch-Godfrey test). In order to minimize residuals' variance relative to the number of lags, we perform univariate multivariate ARCH-LM tests. We report Johansen rank test results in Annex C as well as lags indicated by information criteria on the unrestricted VAR (so that the VECM lag will be equal to unrestricted VAR lag minus one).

[9]One also has to keep in mind that increased estimation uncertainty due to high lag orders can imply large size distortions of the rank test, as pointed out by Gonzalo and Pitarakis (2006) even for the case of no breaks and only one fitted autoregressive lag.

5.4 Long-term restrictions and weak exogeneity testing

VECM estimation is done using an identifying normalization that is not unique (hence, that is not economically meaningful). In order to give economic sense to β parameters, we test restrictions (which is mathematically equivalent to over-identifying the cointegration space). The long-run restrictions are tested with a Wald test with a $\chi^2(J)$ distribution, with J the number of linearly independent restrictions.
With $\beta^*_{K^*-r}$ being a (K^*-r) matrix (the β estimators below the I_r identity matrix), R being a $J \times (K^*-r)r$ -matrix and r being a J-dimensional vector, vec() the vectorization operator, we test the null hypothesis versus the alternative:

$$\text{H0: } R\text{vec}(\beta^*_{K^*-r}) = r \text{ versus H1: } R\text{vec}(\beta^*_{K^*-r}) \neq \text{r} \tag{5}$$

Given linear restriction test results, long-term relation will be estimated again after exclusion of the variables that do not enter in the cointegrating relation. As a basis for long-term analysis resulting from these restrictions, we use the Uncovered Interest Parity (UIP) that links the nominal exchange rate with short-term rate differential between two countries and the empirical Taylor monetary policy rule linking the setting of the nominal interest rate to the inflation gap, the output gap (augmented with the exchange rate). Following notation from Section 4, α (adjustment parameters, or loading coefficients) and γ parameter estimates (slopes of the cointegrating relations) have an asymptotic normal distribution and their t-ratios can be interpreted the usual way.

i) For α estimates, t-ratios allow to test long-term weak exogeneity, conditional on cointegration parameters. A variable is weakly exogenous with respects to a given set of β parameters (i.e. a given long-term correction term) if not modelling the variable in the system does not incur information loss. Weak exogeneity test is of importance for our model since Euro area policy rate is included as an endogenous variable: it can be weakly exogenous for one or more error-correction terms, and its inclusion may lead other variables to be weakly exogenous relative to the β parameters of interest, too.

ii) γ which are trend parameters included in the cointegrating relation (slopes of the cointegrating relations) have an asymptotic normal distribution and thus its t-ratio can be interpreted the usual way. Restrictions on γ are sequentially tested with a joint Wald test per sub period as per Johansen et al. (2000).

5.5 Alternative models

In order to answer all aspects of our research question, we estimate the following alternative VECMs:

i) No breaks model: Baseline model without structural breaks

$$X_t = \{logIP_t, \Delta CPI_t, i^*_t, \Delta e_t, i_t\}' \tag{6}$$

ii) Inflation model: Baseline model augmented with Euro area inflation rate. Addition of European inflation may help us control for the price puzzle effect but also model the double constraint in terms of inflation (domestic and Euro area one).

$$X_t = \{logIP_t, \Delta CPI_t, \Delta CPI^*_t, i^*_t, \Delta e_t, i_t, tE_t\}' \tag{7}$$

6 Results

Our table results (Tables 4 to 6) are grouped by model, with all countries reported in the same table to allow for comparative analysis. Break dates are indicated at the top of each table. These tables gather long-term $\hat{\beta}$, short-term $\hat{\alpha}$ and deterministic $\hat{\gamma}$ estimates as well as their restriction tests as described in Subsection 5.4. Long-term coefficients are presented in their error correction term (ECT) format. Following rank robustness checks as developed in Appendix B, some cointegrated vectors may be excluded and the rank of the VECM revised down. In some other cases, long-term restriction tests led to some cointegration vectors to be restricted: in this case they are not reported in tables but still in error correction terms graphs in Appendices D to F. These graphs represent the error correction term inclusive of the deterministic terms included in the cointegrating relation (tE_t). All cointegration estimates are done with the JMulti software.

Johansen modified rank test results are available in Appendix C. For all models tested, most CESEE countries are best modelled with two-lag VECMs (Czech Republic, Poland, Bulgaria, Croatia). Romania VECMs embed one lag whereas Hungary VECMs embed no lag. With these lag orders, we have been controlling for low order error autocorrelation. On the other hand, heteroscedasticity has been rather well controlled for these lags as tested by univariate or multivariate ARCH-LM tests. Finally, univariate normality tests usually reject the normality hypothesis for policy rates, largely due to a kurtosis effect (same for the joint normality tests).

Regarding the interpretation of the coefficients of the cointegrating relations as long-run elasticities (or semielasticities), we rely on Johansen (2005) who does a counterfactual experiment using changes of current values to impact long-run values and shows that estimated coefficients can be interpreted as long-run elasticities. This analysis is based the dynamics of a VECM, hence relying on a short-run/long-run dichotomy, as opposed to endogenous/exogenous variables[10]. Due to identifying restrictions and variable ordering, we usually find two types of cointegrating relations: the first are long-term relations between domestic and foreign variables, which answers our research question. When Euro area rate is restricted, we find cointegrating relations between domestic variables, which answers the question of the exchange rate importance in the conduct of monetary policy. We analyse both types of cointegrating relations sequentially.

6.1 Baseline model

All estimation results are available in Table 4. For this VECM, Johansen modified rank test does not reject at 5% three cointegrating relations for Bulgaria, Poland and Hungary and two long-term relations for the Czech Republic. Romania and Croatia do not reject one cointegrating relation between variables at 5%. Following rank robustness checks, some long-term relations have been restricted and are not reported (the first ECT for Bulgaria). Rank has been revised for the Czech Republic model (two cointegrating vectors instead of three). Three error correction terms out of thirteen can be considered as unit vectors (second Bulgarian ECT, first Poland ECT and Romania ECT).

[10] Main reasons are firstly, variables have been transformed in logs or are already expressed in percent (necessary condition to interpret coefficients as elasticities or semielasticities, since variables do not have any numeraire anymore); secondly, because cointegrating relation has been identified and finally, because the desired long-run change needed to interpret coefficients as long-run elasticities is a vector (k) in the space orthogonal to betas, and as such, multiplied by adequate Γ, will equal the short-run changes.

Table 4: Baseline VECM; $X_t = \{logIP_t, \Delta CPI_t, i_t^*, \Delta e_t, i_t, tE_t\}'$

	Czech Republic		**Bulgaria**		**Poland**			**Romania**	**Hungary**			**Croatia**
Break date 1	1998 M9		2007 M4		1999 M11			2005 M5	1999 M3			2000 M6
Break date 2	2002 M4		2009:M2		2001 M3			2010 M4	2005 M4			2010 M2
Lags	2		1		2			1	0			1
	ECT1	ECT2	ECT2	ECT3	ECT1	ECT2	ECT3	ECT1	ECT1	ECT2	ECT3	ECT1
β_{IP}	1	*	*	*	1	*	*	1	1	*	*	1
$\beta_{\Delta CPI}$	*	1.00	1.00	*	*	1.00	*	*	*	1.00	*	*
β_{i*}	0.23	*	*	1.00	*	*	1.00	-0.011	*	*	1.00	-0.014
$\beta_{\Delta e}$	0.041	0.12	-	-	0.002	0.07	*	-0.006	0.13	1.87	1.69	-0.004
β_i	-0.380	-0.78	0.74	-0.88	0.02	-0.55	*	0.033	-0.17	-3.61	-2.25	*
γ_{t1}	*	*	0.081**	-0.002	0.043***	1.289***	0.231*	*	-0.04***	-0.431***	-0.377**	*
γ_{t2}	-0.00127	-0.095***	0.073***	0.014***	-0.01**	-0.109*	0.015	*	-0.018**	-0.207	-0.157	-0.00243
γ_{t3}	-0.00230	-0.01**	0.016	0.004	-0.005***	-0.011	0.024***	*	-0.012***	-0.193**	-0.133*	*
α_{IP}	-0.009	**-0.004*** **	0.001	0.003	-0.102	0.002	-0.003	**-0.148*** **	**-0.467*** **	-0.002	**0.039*** **	**-0.33*** **
$\alpha_{\Delta CPI}$	0.234	**-0.101*** **	**-0.13*** **	-0.036	0.743	-0.009	**0.058** *	-2.65	0.097	**-0.077*** **	**0.089** **	**2.48*** **
α_{i*}	0.005	-0.011	**0.011** **	-0.016	**0.412** **	-0.013	**-0.039*** **	**0.569** **	**0.347** **	**0.036*** **	**-0.067*** **	0.48
$\alpha_{\Delta e}$	-1.22	**-0.616*** **	-	-	-4.12	-0.185	0.19	8.25	-2.73	**-0.298** **	**0.456** *	4.57
α_i	**0.297** **	-0.019	-0.016	**0.223*** **	-0.132	**0.106*** **	0.03	1.35	-0.455	-0.003	0.05	**5.96*** **
$H_{l,1}^{\gamma}(r)$	0.93		0.000		0.000			0.874	0.000			0.81
$H_{l,2}^{\gamma}(r)$	0.055		0.000		0.001			0.206	0.000			0
$H_{l,3}^{\gamma}(r)$	0.010		0.046		0.000			0.539	0.000			0.38
$H_0^{\Delta CPI}$	*	*	*	*	*	*	*	**0.39**				**0.73**
H_0^{i*}	0.015	**0.377**	*	*	*	*	*	0.030	*	*	*	0.051
$H_0^{\Delta e}$	0.000	0.021	-	-	0	0.2599	**0.423**	0.000	0	0	0	0.052
H_0^{i}	0.000	0.055	0.040	0.000	0.005	0.0039	**0.267**	0.000				**0.526**
Joint	0.83						0.419	0.62-0.45				0.5235-0.4088

Note: The star * included in cointegrating relations means parameter has been restricted (either identification or long term restrictions). βs are estimated parameters of the restricted cointegrating vectors (or error correction terms - ECT-), with their associated long-run exclusion tests p-values ($H_0^{\Delta CPI}$, H_0^{i*}, $H_0^{\Delta e}$, H_0^{i}). γs are estimated parameters of the restricted linear trend by subperiod, with their associated joint exclusion tests p-values ($H_{l,1}^{\gamma}(r)$, $H_{l,2}^{\gamma}(r)$, $H_{l,3}^{\gamma}(r)$). αs are the loading coefficients defining the adjustment speed of the variables to the ECT (if significant, we reject the null of the variable being weakly exogenous conditional to the ECT). *, **, *** correspond to significance levels at 10%, 5% and 1% respectively.

First interesting result is that **structural break modelling is not rejected at 5% for any of the first EU joiners (namely the Central Eastern countries) and Bulgaria, but is fully rejected for Romania and on two subperiods over three for Croatia**. Estimated $\hat{\gamma}$ coefficients are particularly significant for the three Central countries (Hungary, Poland, and to a lesser extend, the Czech Republic) and with the expected negative sign. This result is in line with the longer data length used for Central Eastern European countries, but rather unexpected from this perspective for Croatia. One can nevertheless argue that this country has a rather stable exchange rate nominal target since 1997, which can explain the rejection of structural trend breaks in the cointegrating relation in a similar fashion as for Romania.

Second interesting result relates to the interdependence between domestic and Euro area monetary policies: **domestic policy rate and other domestic variables are in a long-term relation with Euro area policy rate for three countries of interest: the Czech Republic, Hungary and Bulgaria.** If we normalize the error correction term over the domestic policy rate, we find estimated $\hat{\beta_i *}$ range from 0.45 to 0.6, with the exception of Bulgaria for which we find an almost unitary relation between policy rates ($\hat{\beta}_{CZK} = 0.61, \hat{\beta}_{BUL} = 1.13, \hat{\beta}_{HUF} = 0.44$). More precisely, Euro area and domestic policy rates cointegrate within a Taylor rule augmented with Euro area policy rate (ECT1 for the Czech Republic, ECT3 for Bulgaria) whereas for Hungary, policy rates cointegrate within a UIP type relationship (ECT3). The related $\hat{\alpha}$ are in favor of this interpretation: they have the expected sign[11] and their large magnitude (around 0.3 and even 0.45 for Hungary) and strong significance show that they lead the short-term adjustment of the cointegrating relation. **For these countries as well, the Euro area policy rate i* is weakly exogenous at 5% level conditional on the estimated long-term parameters, still with the exception of Hungary.** For this country, all variables adjust to the long-term disequilibrium (particularly the exchange rate variation) with the exception of the policy rate which is not significant at 5%. On the other hand, **Poland policy rate does not cointegrate with Euro area policy rate** (all domestic variables have been restricted from the third error correction term, leading to a EONIA unit vector). If we consider the second error correction term, we can see that the monetary policy rule is probably well described by a Taylor rule augmented with the exchange rate variation and a strong weight on output but independent from the Euro area monetary policy, even thought identifying restriction on industrial production does not allow us to precisely estimate the coefficients of the Taylor rule. The model also found such a relation for the Czech Republic (ECT2).
Finally, we cannot find a cointegrating relation for Croatia and Romania monetary variables and Euro area policy rates. This is because our model could not manage to identify a cointegrating relation between variables and only modelled a output unit vector (corroborated by the adjustment coefficients that are only significant for the IPI at 5% level). We do not validate this type of modelling for these two countries.

Third interesting result is that **the exchange rate variation is never restricted at 5% level from any cointegrating relation** (with low estimated coefficients for the Czech Republic and Poland, much higher for Hungary). Moreover, they adjust quite largely on the short term for Hungary and the Czech Republic (alphas between 0.45 to

[11] Positive for policy rate since they have negative coefficients in the ECTs, positive for the exchange rate variation in the case of Hungary

-0.62), but the adjustment does not come from the same type of long-term relation. The Czech koruna adjusts from a disequilibrium in monetary policy whereas the Hungarian forint is influenced by international transmission of interest rates. Another interesting result is that exchange rate variation is weakly exogenous at 5% for Poland (it is also the case for Croatia and Bulgaria, but as seen above, we don't validate the modelling for these two countries).

6.2 No breaks model

All estimation results are in Table 5. To be aligned with the baseline model, no breaks VECMs include a restricted linear trend and an unrestricted constant, thus allowing both stationary and non-stationary series to exhibit trends. **For most cointegrating relations though, the inclusion of linear trend is rejected at 5%.** Regarding the number of cointegrating vectors per country, for most countries except Romania and Hungary, Johansen rank test does not reject at 5% two cointegrating relations (instead of three when breaks were included). On the contrary, we cannot reject the null of no cointegrating relation for Croatia. As for the baseline model, we find several unit vectors (four in total), which means the rank test has misidentified the number of cointegrating relations by one. We found four Taylor-type monetary long-term relations (with a restricted Euro area policy rate : Poland ECT2, Hungary ECT2 and without restriction on the Euro area policy rate: Bulgaria ECT2, Romania ECT2).

Regarding the influence of Euro area monetary policy in this model, first striking result is that **Euro area policy rate is long-term restricted to zero for the Czech Republic and for Hungary at 5%**. We therefore cannot find any Euro area monetary spillover from Euro area to these countries, which strongly differs from our baseline VECM results. On the other hand, we still find a long-term relationship between domestic and Euro area policy rates for Bulgaria ($\hat{\beta}_{BUL} = 1.16/0.85$), and still no influence of Euro area policy rate in the cointegrating relations for Poland. We find some significant but small feedback from domestic monetary cointegrating vectors to the Euro area policy rate (Poland ECT2, Hungary ECT2), which was already the case in the baseline model. Overall, the Euro policy rate is weakly exogenous at 5% conditional on long-term parameters of economically-significant cointegrating vectors.
Romania results are interesting because since the structural break modelling was rejected in the baseline model, the no breaks VECM is expected to bring better results. **Romania second cointegrating vector can be interpreted as a domestic monetary policy augmented with the Euro are policy rate and the exchange rate variation**, with a small weight on both and a large weight on domestic inflation (normalized on i, we find $\hat{\beta}_e = 0.10, \hat{\beta}_{i*} = 0.42$ and $\hat{\beta}_{CPI} = 1.20$) and a rather slow adjustment of the domestic rate to a long-term disequilibrium ($\hat{\alpha}_i = 0.059$)[12].
Regarding the exchange rate inclusion in the long-term relations, no-breaks modelling yields the same results as the structural breaks modelling because exchange rates are not long-term restricted to zero as per baseline VECM. We find the same long-term influence of exchange rate in domestic monetary policy for Hungary as in the baseline model ($\hat{\beta}_e = 0.23$) and very strong adjustment to the error correction term.

[12]First error correction term is difficult to interpret, since we find unexpected negative signs on i* and e, but quite significant associated alphas for almost all variables.

Table 5: VECM without structural breaks; $X_t = \{logIP_t, \Delta CPI_t, i^*, \Delta e_t, i_t, t\}'$

	Czech Republic	**Bulgaria**		**Poland**		**Romania**		**Hungary**		
Break date 1	1998 M9	2007 M4		1999 M11		2005 M5		1999 M3		
Break date 2	2002 M4	2009:M2		2001 M3		2010 M4		2005 M4		
Lags	2	2		2		1	2	0		
	ECT1	ECT1	ECT2	ECT1	ECT2	ECT1	ECT2	ECT1	ECT2	ECT3
β_{IP}	1	1	*	1.000	*	1.000	*	1.000	*	*
$\beta_{\Delta CPI}$	-0.074	*	1.00	*	1.000	*	1.000	*	1.000	*
β_{i*}	*	0.19	2.549	-0.038	*	-0.044	0.353	*	*	1.000
$\beta_{\Delta e}$	-0.005	-	-	0.0040	0.044	-0.003	0.088	0.018	0.252	0.209
β_i	0.084	-0.15	-2.987	-0.001	-0.623	-0.006	-0.83	0.006	-1.09	*
γ_{t1}	*	*	0.038**	-0.007***	*	-0.005***	*	-0.002***	-0.038***	*
α_{IP}	-0.019	-0.023*	-0.001	-0.019	0.000	**-0.228****	-0.001	**-0.32*****	-0.004	**0.027*****
$\alpha_{\Delta CPI}$	**1.75****	**2.01*****	**-0.121*****	**2.22*****	**-0.059*****	2.06	**-0.09*****	**0.659****	0.015	-0.047
α_{i*}	**0.543*****	**0.221***	0.004	**0.584*****	**-0.017***	**1.63*****	0.001	**0.27*****	**0.05*****	**-0.073*****
$\alpha_{\Delta e}$	-0.84	-	-	**-13.63****	0.210	**7.82**	**0.245***	-3.19	**-0.632*****	**0.67***
α_i	0.60	0.24	0.006	0.062	**0.066*****	**4.3*****	**0.059*****	0.22	0.02	**0.09**
$H_l^{\gamma}(r)$	0.838	0.834	0.004	0.00	0.37	0.000	0.62	0.307	0.018	0.54
$H_0^{\Delta CPI}$	0.00	*	*	*	*	*	*	*	*	*
H_0^{i*}	**0.61**	0.000	0.000	0.139	**0.67**	0.000	0.112	*	*	*
$H_0^{\Delta e}$	0.00	-	-	0.000	0.000	0.000	0.000	0.000	0.000	0.000
H_0^{i}	0.00	0.000	0.000	0.046	0.000	0.028	0.030	0.45	0.00	0.84
Joint	0.680	0.006		0.043	0.32		0.05	0.018	0.00	0.72

Note: The star * included in cointegrating relations means parameter has been restricted (either identification or long term restrictions). βs are estimated parameters of the restricted cointegrating vectors (or error correction terms - ECT-), with their associated long-run exclusion tests p-values (H_0^{i*}, $H_0^{\Delta e}$, H_0^{i}). γs are estimated parameters of the restricted linear trend by subperiod, with their associated joint exclusion tests p-values ($H_{l,1}^{\gamma}(r)$, $H_{l,2}^{\gamma}(r)$, $H_{l,3}^{\gamma}(r)$). αs are the loading coefficients defining the adjustment speed of the variables to the ECT (if significant, we reject the null of the variable being weakly exogenous conditional to the ECT). *, **, *** correspond to significance levels at 10%, 5% and 1% respectively.

6.3 Inflation model

The inclusion of Euro area inflation rate in our baseline VECM aims at testing whether the Maastricht inflation criterion (as highlighted in Section 2) has played a long-term role in the conduct of monetary policy of CESEE countries. If we first take a look at the number of cointegrating relations now found by the modified Johansen rank test in Annex C, we can observe that **the inclusion of Euro area inflation rate in the system has reduced the number of cointegrating relations by one for Poland and Hungary**. The other countries are not impacted in this respect by the inclusion of Euro area inflation. All estimation results are available in Table 6[13]. We have not estimated Romania model because Euro area inflation is stationary for the data sample considered (as shown in Table 2). After having initially estimated Bulgaria and Czech Republic VECMs with three cointegrating relations, we have finally re-estimated the VECMs with a rank of two, following Annex B rank robustness checks procedure.

Structural break modelling is overall validated, with nevertheless some sub-period restrictions to zero at 5% level compared to the baseline model defined in Equation 4 (particularly for the Czech Republic). Estimated trends are lower than in Baseline VECMs.

First striking result is that **Euro area inflation is not restricted at 5% from the long-run relationships between domestic variables for the Czech Republic and Bulgaria.** Nevertheless, it is long-term fully restricted and weakly exogenous at 5% relative to long-term parameters for Poland and Hungary. **Euro area policy rate impact is robust to the inclusion of Euro inflation for Bulgaria only**: the variable is not restricted and enters the cointegrating relation (β_{i*}=0.79). It is restricted on the long-term for Poland, Croatia and has a wrong sign in Hungary[14] and the Czech Republic cointegrating vectors. Finally, as mentioned above, the modelling of Croatia is highly perturbed by the inclusion of Euro area inflation, since the model ends up modelling a unit vector. We therefore believe that **the inclusion of Euro Area inflation in the Baseline VECM is not justified for Poland, Croatia and Hungary**.

Second interesting result is that we find a **long-term relation between Euro area inflation and domestic inflation for the Czech Republic and Bulgaria, in the form of a inflation differential** (both in their respective ECT2), with $\beta_{\Delta CPI*}$= -1.52 and -0.78 respectively[15]. Noticeably in the case of the Czech Republic, the exchange rate variation very strongly adjusts to this long-run equilibrium relation. Finally for Hungary, we find a long-term relationship in the form of a monetary policy rule in ECT2, with a strong weight on domestic inflation as in the no breaks VECM. For Hungary and the Czech Republic, the estimated coefficient on domestic inflation (when normalized over domestic policy rate) verifies the Taylor condition.

[13] One can see that five cointegrating vectors out of nine can be considered as unit vectors - all first cointegrating vectors for each country

[14] We did not restrict Euro area policy rate due to joint Wald test results

[15] The domestic policy rate is weakly exogenous at 5% for these two error-correction terms.

Table 6: Inflation VECM; $X_t = \{logIP_t, \Delta CPI_t, \Delta CPI_t^*, i^*, \Delta e_t, i_t, tE_t\}'$

	Czech Republic		Bulgaria		Poland		Hungary		Croatia
Break dates	1998 M9 2002 M4		2007 M4 2009:M2		1999 M11 2001 M3		1999 M3 2005 M4		2000 M6 2010 M2
Lags	2		2		2		2		2
	ECT1	ECT2	ECT1	ECT2	ECT1	ECT2	ECT1	ECT2	ECT1
β_{IP}	1	*	1	*	1	*	1	*	1
$\beta_{\Delta CPI}$	*	1.00	*	1.00	*	1.00	*	1.00	0.01
$\beta_{\Delta CPI*}$	*	-1.24	0.003	-2.344	*	*	*	*	-0.029
β_{i*}	-0.05	-0.39	-0.007	2.354	-0.02	0.28	-0.06	*	-0.02
$\beta_{\Delta e}$	0.00	*	-	-	0.00	0.07	0.00	-0.04	0.00
β_i	-0.01	-0.81	0.022	-2.971	0.000	*	0.00	-1.01	*
t1	*	*	-0.006***	-0.008	0.03***	1.31***	-0.013***	0.07***	*
t2	*	*	*	*	-0.013***	0.11	-0.006***	0.01	-0.003***
t3	-0.003***	-0.022***	-0.0004	0.036**	-0.006***	0.016*	0.000	-0.004	*
α_{IP}	-0.023	-0.001	-0.254***	0.0006	-0.04	-0.001	-0.67**	-0.01***	-0.457***
$\alpha_{\Delta CPI}$	2.45***	-0.169***	0.8	-0.139***	2.65***	-0.049**	-0.35	-0.11***	1.06
$\alpha_{\Delta CPI*}$	1.45***	-0.02	-0.7	0.0411***	1.23***	-0.005	-0.03	0.025***	1.2*
α_{i*}	0.98***	-0.018	0.5	0.019***	1.02***	-0.024***	0.33	0.01**	0.55
$\alpha_{\Delta e}$	-5.7	-0.958***	-	-	-18.3***	-0.104	-1.66	-0.228	6.6*
α_i	0.6	-0.02	0.8	0.016	0.81	0.084***	0.18	0.028	7.07***
$H^{\gamma}_{l,1}(r)$	0.43		0.0001		0		0.00		0.30
$H^{\gamma}_{l,2}(r)$	0.82		0.0868		0.00		0.00		0.0001
$H^{\gamma}_{l,3}(r)$	0.000		0.0265		0.00		0.00		0.97
$H_0^{\Delta CPI}$	*	*	*	*	*	*	*	*	0.019
$H_0^{\Delta CPI*}$	0.51	0	0.019	0	0.26	0.91	0.28	0.67	0
H_0^{i*}	0.015	0.006	0.000	0.000	0.03	0.14	0.00	0.05	0.057
$H_0^{\Delta e}$	0.000	0.749	-	-	0.000	0.00	0.00	0.00	0.10
H_0^{i}	0.000	0.44	0.000	0.000	0.06	0.866	0.000	0.017	0.21
H0 jointes	0.75		NR		i, CPI*=0.72		0.08		0.52

Note: The star * included in cointegrating relations means parameter has been restricted (either identification or long term restrictions). βs are estimated parameters of the restricted cointegrating vectors (or error correction terms - ECT-), with their associated long-run exclusion tests p-values ($H_0^{\Delta CPI}$, $H_0^{\Delta CPI*}$ H_0^{i*}, $H_0^{\Delta e}$, H_0^{i}). γs are estimated parameters of the restricted linear trend by subperiod, with their associated joint exclusion tests p-values ($H^{\gamma}_{l,1}(r)$, $H^{\gamma}_{l,2}(r)$, $H^{\gamma}_{l,3}(r)$). αs are the loading coefficients defining the adjustment speed of the variables to the ECT (if significant, we reject the null of the variable being weakly exogenous conditional to the ECT). *, **, *** correspond to significance levels at 10%, 5% and 1% respectively.

6.4 Model discussion and specific robustness checks

For baseline and inflation VECMs, we have found that structural break modelling is indeed significant for early EU joiners (Central Eastern European countries) and Bulgaria. This is quite interesting because we chose to use invariant structural break dates that despite being determined endogenously, were in relation with countries' history as discussed in Subsection 3.2. On the contrary, this type of modelling is not suited for Romania and Croatia, even though we include a structural break date that is not domestic and more potentially linked to the GFC (beginning of 2010 for both countries).

In several inflation and no breaks VECMs, the Johansen modified rank test has identified one cointegrating relation in excess, as can be inferred when Industrial Production unit vector is modelled as the first cointegrating vector. We think that the parametric nature of both the unit root and the rank tests (i.e. lags used to control for series autocorrelation in the unit root test and to whiten the residuals in the unrestricted VAR for the modified rank test) has to be counterbalanced with the rank robustness checks, as described in Appendix B and as we have performed throughout this paper. The parametric sensitivity is however inherent to VAR and VECM modelling and is not specific to our structural break modelling.

We are specifically interested in assessing the sensitivity of the Clemente et al. (1998) unit root test, particularly as a robustness check for endogenous break dates determination. We have compared the two types of innovations and two lag trucation methods for the Clemente et al. (1998) unit root test for the Industrial Production series. It is the only series in levels with policy rates, which allows us to test unit root test robustness on first differenced series as well. We use both the Additional Outlier (AO) and the Innovative Outlier (IO) methods. For the robustness of the lag truncation method, we use the partial autocorrelation function of the second-differenced series and input the new lag in the unit root algorithm. With L the lag operator, the second-differenced series is calculated as follow: (1-L)$\Delta X_t = X_t$ - $2X_{t-1}$ + X_{t-2}.

From this example in Table 8, we can infer that:

i/ The unit root test on series in levels is robust to any type of innovation model when series display high level of partial autocorrelation. Both models roughly calculate the same endogenous break dates and lags.

ii/ The unit root test with IO model better captures the global endogenous break dates (such as the one related to the GFC) in first-differenced series, even when the number of lags to control series autocorrelation is low (see for instance the results for Romania, Bulgaria and Croatia). This results in the fact that we reject the null hypothesis of the presence of a unit root with possibly one break in mean more often when we use an Innovative Outlier model for first-differenced series.

iii/ The Perron and Vogelsang (1992) lag truncation method is largely robust for the determination of lags if we compare with the lags obtained from the partial autocorrelation analysis on second-differenced series.

iiii/ Finally, we can see the impact of lag parameters in the unit root test results.

7 Conclusion and further research

Testing monetary autonomy of CESEE countries in the context of EMU convergence using multivariate cointegration with structural breaks has revealed conventional monetary interdependence between the Czech Republic, Bulgaria, Hungary and Romania and the Euro area. These interdependences are rather aligned in terms of amplitude with the exchange rate regime of CESEE countries: the less flexible the exchange rate regime (or the higher the degree of foreign exchange interventions, as for the Czech Republic for instance), the higher the monetary spillover from Euro area to CESEE countries. Nevertheless, this spillover is seldom identified as coming from a UIP relationship: this result is interesting and validates the multivariate approach. One exception lies with Croatia, which maintains a soft peg with Euro and for which we have not managed to model any monetary relation with the Euro area, even in a no structural break framework. Euro area policy rate and inflation are in most cases weakly exogenous conditional on the long-term parameters: they are part of the long-term relations but do not adjust to them.

One of the main difficulties to overcome to proceed with this long-term analysis has been to correctly model the structural breaks inherent to the macroeconomic series of these countries. Structural break modelling has been appropriate for Central and Eastern European countries and Bulgaria; but it is not suited as such for Croatia and Romania. One important implication for the Czech Republic and Hungary is that without structural break modelling, we don't find significant and large long-run Euro area policy rate spillover effect as we find in the structural breaks model. In terms of macroeconomic policy recommendation, our results tend to show that the Czech Republic and Romania could lose both the exchange rate and the monetary policy rate instruments probably at a lower cost compared to Poland. The importance of Euro area inflation pass-through has also been revealed for the Czech Republic and Bulgaria, which also points towards nominal convergence for the former country. In Hungary, Euro area policy rate transmission is driven by the international transmission of interest rates through the exchange rate. We cannot conclude for Croatia on this perspective.

Further research could include the Quantitative Easing instruments to see if the new types of monetary policies implemented by the Euro area have had an impact on the degree of monetary autonomy for our countries of interest. We plan to use Shadow Short Rates (SSRs) as calculated by (Krippner, 2015), that account for unconventional monetary policies and are not constrained by the Zero Lower Bound (ZLB) imposed by non-arbitrage conditions on nominal policy rates. As a first extension, we want to use Euro area SSR as a robustness measure of Euro area overall monetary policy stance over our initial time span. As a second step, we need to review the unconventional policy transmission channels identified in the recent literature, extend our data spam (from January 2015 onwards) and calculate CESEE countries' SSRs.

Bibliography

Aizenman, Joshua, Michael Hutchison, and Ilan Noy (2011), "Inflation Targeting and Real Exchange Rates in Emerging Markets." *World Development*, 39, 712–724.

Allegret, Jean Pierre (2012), "Responses of Monetary Authorities in Emerging Economies to International Financial Crises : What Do We Really Know ?" *European Research Studies*, XV, 3–32.

Ball, Laurence (2000), "Policy Rules and External Shocks." *NBER Working Paper Series*, No. 7910.

Bluwstein, Kristina and Fabio Canova (2016), "Beggar-thy-neighbor? The international effects of ECB unconventional monetary policy measures." *International Journal of Central Banking*, 12, 69–120.

Brada, Josef C., Ali M. Kutan, and Su Zhou (2005), "Real and monetary convergence between the European Union's core and recent member countries: A rolling cointegration approach." *Journal of Banking and Finance*, 29, 249–270.

Calvo, Guillermo A and Carmen M Reinhart (2002), "Fear of Floating." *Quarterly Journal of Economics*, CXVII, 379–408.

Camarero, Mariam, Javier Ordón Ez, and Cecilio R. Tamarit (2002), "Tests for interest rate convergence and structural breaks in the EMS : further analysis." *Applied Financial Economics*, 12, 447–456.

Clemente, Jesús, Antonio Montañés, and Marcelo Reyes (1998), "Testing for a unit root in variables with a double change in the mean." *Economics Letters*, 59, 175–182.

Creel, Jérôme and Sandrine Levasseur (2006), "Canaux de transmission de la politique monétaire dans l'UE." *Revue économique*, 57, 881–898.

Crespo-Cuaresma, Jesús, Haralad Oberhofer, Karlis Smits, and Gallina A Vincelette (2012), "Drivers of Convergence in Eleven Eastern European Countries." *World Bank Policy Research Working Paper*, 6185.

Crespo-Cuaresma, Jesús and Cezary Wójcik (2006), "Measuring monetary independence: Evidence from a group of new EU member countries." *Journal of Comparative Economics*, 34, 24–43.

Căpraru, Bogdan and Iulian Ihnatov (2012), "Interest Rate Transmission and Exchange Rate Arrangements in the Central and Eastern European Countries: Evidence from the Current International Financial Crises." *Procedia - Social and Behavioral Sciences*, 58, 1273–1282.

Égert, Balázs and Ronald Macdonald (2009), "Monetary transmission mechanism in Central and Eastern Europe: Surveying the surveyable." *Journal of Economic Surveys*, 23, 277–327.

Elbourne, Adam and Jakob de Haan (2009), "Modeling Monetary Policy Transmission in Acceding Countries: Vector Autoregression Versus Structural Vector Autoregression." *Emerging Markets Finance and Trade*, 45, 4–20.

European Central Bank (2020), *Convergence Report*, volume 8014.

European Commission (2018), “Convergence Report.” Technical report, Directorate-General for Economic and Financial Affairs.

Fountas, Stilianos and Jyh-lin Wu (1998), “Tests for interest rate convergence and structural breaks in the EMS.” *Applied Financial Economics*, 127–132.

Frankel, Jeffrey A., Sergio L. Schmukler, and Luis Servén (2004), “Global transmission of interest rates: Monetary independence and currency regime.” *Journal of International Money and Finance*, 23, 701–733.

Frömmel, Michael, Garo Garabedian, and Franziska Schobert (2011), “Monetary policy rules in Central and Eastern European Countries: Does the exchange rate matter?” *Journal of Macroeconomics*, 33, 807–818.

Frömmel, Michael and Robinson Kruse (2015), “Interest rate convergence in the EMS prior to European Monetary Union.” *Journal of Policy Modeling*, 37, 990–1004.

Gallego, Sonsoles, Sándor Gardó, Reiner Martin, Luis Molina, and Jose María Serena (2010), “The Impact of the Global Economic and Financial Crisis on Central, Eastern and South-Eastern Europe (CESEE) and Latin America.” *Banco de Espana Documentos Ocasionales*, 1002.

Gerlach-Kristen, Petra (2003), “Interest rate reaction functions and the Taylor rule in the Euro area.” *European Central Bank Working Paper Series*, 258.

Ghosh, Atish R., Jonathan D. Ostry, and Marcos Chamon (2016), “Two targets, two instruments: Monetary and exchange rate policies in emerging market economies.” *Journal of International Money and Finance*, 60, 172–196.

Gonzalo, Jesús and Jean Yves Pitarakis (2006), “Threshold effects in cointegrating relationships.” *Oxford Bulletin of Economics and Statistics*, 68, 813–833.

Gregory, Allan W. and Bruce E. Hansen (1996), “Residual-based tests for cointegration in models with regime shifts.” *Journal of Econometrics*, 70, 99–126.

Habib, Maurizio Michael (2002), “Financial contagion, interest rates and the role of the exchange rate as shock absorber in Central and Eastern Europe.” *BOFIT Discussion Papers*, 7.

Hajek, Jan and Roman Horvath (2018), “International spillovers of (un)conventional monetary policy: The effect of the ECB and the US Fed on non-euro EU countries.” *Economic Systems*, 42, 91–105.

Hansen, Bruce E. and Byeongseon Seo (2002), “Testing for Threshold Cointegration in Vector Error Correction Models.” *Journal of Econometrics*, 110, 293–318.

Hendry, David F. and Katarina Juselius (2001), “Explaining Cointegration Analysis: Part 2.” *Energy Journal*, 22, 75–120.

Holtemöller, Oliver (2005), “Uncovered interest rate parity and analysis of monetary convergence of potential EMU accession countries.” *International Economics and Economic Policy*, 2, 33–63.

Ishi, Kotaro, Mark Stone, and Etienne B Yehoue (2009), "Unconventional Central Bank Measures for Emerging Economies." *IMF Working Paper Series*, 226.

Jiménez-Rodríguez, Rebeca, Amalia Morales-Zumaquero, and Balázs Égert (2010), "The effect of foreign shocks in Central and Eastern Europe." *Journal of Policy Modeling*, 32, 461–477.

Johansen, Søren (1995), *Likelihood-Based Inference in Cointegrated Vector Autoregressive Models*. Oxford University Press.

Johansen, Søren (2005), "Interpretation of Cointegrating Coefficients in the Cointegrated Vector Autoregressive Model." *Oxford Bulletin of Economics and Statistics*, 67, 0305–9049.

Johansen, Søren, Rocco Mosconi, and Bent Nielsen (2000), "Cointegration analysis in the presence of structural breaks in the deterministic trend." *Econometrics Journal*, 3, pp. 216–49.

Josifidis, Kosta, Jean Pierre Allegret, Céline Gimet, and Emilija Beker Pucar (2014), "Macroeconomic policy responses to financial crises in emerging European economies." *Economic Modelling*, 36, 577–591.

Kasman, Adnan, Saadet Kirbas-Kasman, and Evrim Turgutlu (2008), "Monetary policy convergence of potential EMU accession countries: A cointegration analysis with shifting regimes." *Economic Modelling*, 25, 340–350.

Koukouritakis, Minoas, Athanasios P. Papadopoulos, and Andreas Yannopoulos (2014), "Transmission effects in the presence of structural breaks: Evidence from South-Eastern European countries." *Economic Modelling*, 41, 298–311.

Krippner, L. (2015), "A comment on Wu and Xia (2015), and the case for two-factor Shadow Short Rates." *CAMA Working Paper*, 48.

Lee, Junsoo and Mark C . Strazicich (2003), "Minimum Lagrange Multiplier Unit Root Test with Two Structural Breaks." *The Review of Economics and Statistics*, 85, 1082–1089.

Lumsdaine, Robin L. and David H Papell (1997), "Multiple trend breaks and the unit-root hypothesis." *Review of Economics and Statistics*, 79.

Lütkepohl, Helmut (2004), "Vector Autoregressive and Vector Error Correction Models." In *Applied Time Series Econometrics* (Helmut Lütkepohl and Markus Krätzig, eds.), 86–158, Cambridge University Press.

Lütkepohl, Helmut, Pentti Saikkonen, and Carsten Trenkler (2004), "Testing for the cointegrating rank of a var process with level shift at unknown time." *Econometrica*, 72, 647–662.

Magyar Nemzeti Bank (2018), "Future Strategic Framework For The Set Of Unconventional Monetary Policy Instruments Affecting Short-Term Yields."

Magyar Nemzeti Bank (2020), "Monetary Policy Instruments Of The Magyar Nemzeti Bank During The Covid 19 Crisis: Liquidity, Safety, Flexibility."

Matei, Iuliana and Jerôme Héricourt (2006), “Transmission de la politique monétaire dans les pays d ’ Europe centrale et orientale : que savons-nous vraiment ?” *Economie & prévision*, 4-5, 221–238.

Moder, Isabella (2017), “Spillovers from the ECB’ s non-standard monetary policy measures on south-eastern Europe.” *European Central Bank Working Paper Series*, 2095.

Mundell, Robert A (1961), “A Theory of Optimum Currency Areas.” *American Economic Review*, 51, 657–665.

Nelson, Charles R. and Charles R. Plosser (1982), “Trends and random walks in macroeconmic time series. Some evidence and implications.” *Journal of Monetary Economics*, 10, 139–162.

Obstfeld, Maurice, Jay C. Shambaugh, and Alan M. Taylor (2004), “The Trilemma In History: Tradeoffs Among Exchange Rates, Monetary Policies, And Capital Mobility.” *NBER Working Paper Series*, 10396.

Peersman, Gert and Benoit Mojon (2001), “A Var Description Of The Effects Of Monetary Policy In The Individual New Countries Of The Euro Area.” *ECB Working Paper Series*, 92.

Peersman, Gert and Frank Smets (2001), “The Monetary Transmission Mechanis in the Euro Area: More Evidence from VAR Analysis.” *ECB Working paper*, 91.

Perron, Pierre (1989), “The Great Crash , the Oil Price Shock , and the Unit Root Hypothesis.” *Econometrica*, 57, 1361–1401.

Perron, Pierre and Timothy J. Vogelsang (1992), “Nonstationarity and level shifts with an application to purchasing power parity.” *Journal of Business and Economic Statistics*, 10, 301–320.

Rey, Hélène (2015), “Dilemma not trilemma: The global financial cycle and monetary policy independence.” *NBER Working Paper Series*, 21162.

Saikkonen, Pentti and Helmut Lütkepohl (2000), “Testing for the Cointegrating Rank of a VAR Process with an Intercept.” *Econometric Theory*, 16, 373–406.

Taylor, John B. (2001), “The role of the exchange rate in monetary policy rules.” *American Economic Review*, 91, 263–267.

Trenkler, Carsten, Pentti Saikkonen, and Helmut Lütkepohl (2008), “Testing for the cointegrating rank of a VAR process with level shift and trend break.” *Journal of Time Series Analysis*, 29, 331–358.

Windberger, Thomas, Jesús Crespo-Cuaresma, and Janette Walde (2012), “Dirty floating and monetary independence in Central and Eastern Europe - The role of structural breaks.” *Working Papers in Economics and Statistics, University of Innsbruck*, 21.

Wollmershäuser, Timo (2006), “Should central banks react to exchange rate movements? An analysis of the robustness of simple policy rules under exchange rate uncertainty.” *Journal of Macroeconomics*, 28, 493–519.

Zivot, Eric and Donald W.K. Andrews (1992), “Further evidence on the great crash, the oil-price shock, and the unit-root hypothesis.” *Journal of Business and Economic Statistics*, 10, 251–270.

Table 7: Data information

Variable	**Definition**	**Description**	**Source**
IPI_t	Industrial Production Index at time t	Total Index; Monthly, end of period; NSA	IMF IFS
CPI_t	National Consumer Price Index at time t	Total Index; Monthly, end of period; NSA; 1995=100	National Statistics Offices
CPI_t	Harmonized Consumer Price Index at time t	All-Items ; Monthly, end of period; NSA; 2005=100	Eurostat
i_t	Central Bank policy rate at time t		National Central Banks
i*_t	Euro area policy rate	Monthly fixing, end of period - Proxied by EONIA	OECD
e	Nominal exchange rate against Euro	Monthly fixing	European Central Bank
log_IPI_t	Natural log of Industrial Production Index	Seasonal adjustement (X-12 ARIMA) Log transformation	Author's calculations
Varia_CPI_t	National inflation rate (annual variation of National Consumer Price Index)	Annual log return (log CPI_t - log CPI_(t-12))	Author's calculations
Varia_HCPI_t	National Harmonized inflation rate (annual variation of Harmonized Consumer Price Index)	Seasonal adjustement Annual log return: (log HCPI_t - log HCPI_(t-12))	Author's calculations
Varia_e_t	Annual variation of exchange rate	Annual log return: (log HCPI_t - log HCPI_(t-12))	Author's calculations

Appendix A Clemente et al. (1998) unit root test

The unit root test tests the null hypothesis of a unit root with possibly one endogenous break in mean, against the alternative of no unit root with two endogenous breaks.

$$\begin{aligned} H0: y_t &= y_{t-1} + \delta_1 DTB_{1t} + \delta_2 DTB_{2t} + u_t \\ H1: y_t &= \mu + d_1 DU_{1t} + d_2 DU_{2t} + e_t \end{aligned} \tag{8}$$

There is a two-step testing procedure for the Additive Outlier case: first estimate an auxiliary regression to get rid of the deterministic component of H1, which is:

$$y_t = \mu + d_1 DU_{1t} + d_2 DU_{2t} + \tilde{y}_t \tag{9}$$

then search for the minimum t-ratio for ρ being closest to the null hypothesis using the second specification:

$$\tilde{y}_t = \sum_{0<i<k} \omega_{1t}\delta_1 DTB_{1t-i} + \sum_{0<i<k} \omega_{2t}\delta_1 DTB_{2t-i} + \rho \tilde{y_{t-1}} + \sum_{0<i<k} c_i \Delta \tilde{y_{t-1}} + e_t \tag{10}$$

Table 8: Robustness of unit root test

Country	Variable	Model	Lag trucation method	Lag truncation	Optimal break point 1	Optimal break point 2	min T-Stat
Poland	IPI	AO	Perron and Vogelsang (1992)	1	2003:M9	2009:M12	-3.71
	IPI	IO	Perron and Vogelsang (1992)	24	2003:M2	2010:M1	-3.57
	ΔIPI	AO	Perron and Vogelsang (1992)	8	1999:M10	2002:M5	-3.49
	ΔIPI	AO	Partial Auto Correlation	6	NA	NA	NA
	ΔIPI	IO	Perron and Vogelsang (1992)	8	2008:M1	2009:M1	-4.11
Czech Republic	IPI	AO	Perron and Vogelsang (1992)	24	2003:M5	2014:M6	-3.21
	IPI	IO	Perron and Vogelsang (1992)	22	2003:M6	2013:M6	-2.58
	ΔIPI	AO	Perron and Vogelsang (1992)	10	2008:M4	2009:M1	-6.19
	ΔIPI	AO	Partial Auto Correlation	1	2008:M4	2009:M11	-16.75
	ΔIPI	IO	Perron and Vogelsang (1992)	10	2008:M5	2008:M12	**-7.45**
Hungary	IPI	AO	Perron and Vogelsang (1992)	22	1998:M12	2003:M6	-2.63
	IPI	IO	Perron and Vogelsang (1992)	21	1999:M1	2003:M7	-2.74
	ΔIPI	AO	Perron and Vogelsang (1992)	12	2008:M10	2009:M11	-3.38
	ΔIPI	AO	Partial Auto Correlation	9	2008:M10	2009:M11	**-8.05**
	ΔIPI	IO	Perron and Vogelsang (1992)	12	2008:M11	2009:M12	-3.76
Romania	IPI	AO	Perron and Vogelsang (1992)	12	2006:M6	2012:M4	-2.77
	IPI	IO	Perron and Vogelsang (1992)	22	2005:M6	2010:M7	-2.00
	ΔIPI	AO	Perron and Vogelsang (1992)	9	2008:M8	2009:M1	-4.91
	ΔIPI	AO	Partial Auto Correlation	6	NA	NA	NA
	ΔIPI	IO	Perron and Vogelsang (1992)	9	2008:M8	2008:M12	**-7.12**
Bulgaria	IPI	AO	Perron and Vogelsang (1992)	5	2004:M12	2008:M7	-2.54
	IPI	IO	Perron and Vogelsang (1992)	7	2002:M2	2010:M1	-3.79
	ΔIPI	AO	Perron and Vogelsang (1992)	11	2008:M7	2009:M3	-4.06
	ΔIPI	AO	Partial Auto Correlation	5	2008:M7	2009:M3	**-7.40**
	ΔIPI	IO	Perron and Vogelsang (1992)	11	2008:M4	2008:M12	-4.83
Croatia	IPI	AO	Perron and Vogelsang (1992)	4	2003:M11	2010:M7	-2.52
	IPI	IO	Perron and Vogelsang (1992)	23	2002:M5	2011:M5	-3.02
	ΔIPI	AO	Perron and Vogelsang (1992)	9	2000:M11	2008:M3	-5.64
	ΔIPI	AO	Partial Auto Correlation	10	NA	NA	NA
	ΔIPI	IO	Perron and Vogelsang (1992)	7	2000:M12	2008:M12	**-8.43**

Critical values for k(t) at 5%= -5.49 ; at 10%= -5.24
NA: Not available because the algorithm cannot be modified precisely for such lag

Appendix B Alternative rank tests to Johansen et al. (2000) and rank robustness checks

Saikkonen and Lütkepohl (2000) have studied the Johansen et al. (2000) model and show that other tests may be advantageous in terms of local power if there is just a level shift. The modified Johansen test may suffer from a loss of power when using long lag lengths. It has also a tendency for size distortions, with over-rejection of a correct null hypothesis for a small number of observations of Data Generating Processes with one cointegrating relation. More precisely, two types of cointegration tests have been developed subsequently to this model:
i/ An alternative test with known break dates and deterministic trend breaks, both in slope and level, developed by Trenkler et al. (2008). The deterministic terms are included in the data generating process directly and not into the VECM representation. There is therefore detrending before rank testing.
ii/ Other tests with unknown break dates and a shift in level such as in Lütkepohl et al. (2004)

We follow the rank robustness check procedure explained in Hendry and Juselius (2001). Firstly, we test several lag lengths for the rank test. Secondly, the rank test may not discriminate near integrated from stationary processes, so that we can end up with modelling near integrated variables in the VECM instead of stationary long-run relationships. To avoid this, on top of the rank test results, one therefore needs to check the following points in order to correctly determine the cointegration rank of the system (which is done on the unrestricted VECM):
-The t-values of the alpha coefficients: for the r^{th+1} cointegrating vector, if they are small (below 3), then there is not much gain in including this vector as a cointegrating relation in the model.
- The recursive graph of the trace statistic: it should increase linearly for the first r components but stay constant for the remainder.
- The graphs of the cointegrating relations $\beta_i' Xt$: if the graphs reveal non-stationary behavior of a cointegration relation, then the rank choice has to be reconsidered or the model specification reassessed.
- The economic interpretability of the results.

Appendix C Johansen modified rank test results

Croatia

H0 : r	p-value	H0 : r	p-value	H0 : r	p-value
0	0.0139	**0**	**0.379**	0	0.0192
1	**0.1407**	1	0.698	**1**	**0.168**
2	0.5799	2	0.722	2	0.434
3	0.4954	3	0.692	3	0.653
4	0.5441	4	0.540	4	0.607
				5	0.756
Lags	**3**	Lags	3	Lags	**3**
Lags AIC	3	Lags AIC	3	Lags AIC	3
Lags HQ	2	Lags HQ	3	Lags HQ	1
Lags SIC	1	Lags SIC	2	Lags SIC	1

Table 9: Johansen (2001) modified rank test results - continued

VECM 4 [logIP,CPI,i*,et,i,tEt]		VECM 6 [logIP,CPI,i*,et,i]		VECM 7 [logIP,CPI,CPI*,i*,et,i,tEt]	
Czech Republic					
H0 : r	p-value	H0 : r	p-value	H0 : r	p-value
0	0.000	0	0.028	0	0.000
1	0.005	**1**	**0.208**	1	0.001
2	0.035	2	0.794	2	0.017
3	**0.271**	3	0.8027	**3**	**0.080**
4	0.4806	4	0.9279	4	0.515
				5	0.937
Lags	3	Lags	3	Lags	3
Lags AIC	3	Lags AIC	3	Lags AIC	3
Lags HQ	2	Lags HQ	2	Lags HQ	1
Lags SIC	1	Lags SIC	1	Lags SIC	1
Romania					
H0 : r	p-value	H0 : r	p-value	H0 : r	p-value
0	0.016	0	0.001	0	0.047
1	**0.297**	1	0.037	**1**	**0.584**
2	0.7896	**2**	**0.3081**	2	0.9044
3	0.7518	3	0.3755	3	0.8642
4	0.6759	4	0.3242	4	0.7711
				5	0.6574
Lags	2	Lags	2	Lags	2
Lags AIC	2	Lags AIC	2	Lags AIC	10
Lags HQ	1	Lags HQ	1	Lags HQ	1
Lags SIC	1	Lags SIC	1	Lags SIC	1
Poland					
H0 : r	p-value	H0 : r	p-value	H0 : r	p-value
0	0.000	0	0.001	0	0
1	0.001	1	0.050	1	0.0008
2	0.046	**2**	**0.052**	**2**	**0.0704**
3	**0.140**	3	0.068	3	0.2863
4	0.175	4	0.069	4	0.5576
				5	0.7779
Lags (levels)	3	Lags (levels)	3	Lags (levels)	3
Lags AIC	3	Lags AIC	3	Lags AIC	10
Lags HQ	3	Lags HQ	3	Lags HQ	3
Lags SIC	2	Lags SIC	2	Lags SIC	1
Bulgaria					
H0 : r	p-value	H0 : r	p-value	H0 : r	p-value
0	0	0	0	0	0
1	0.005	1	0.024	1	0.000
2	0.032	**2**	**0.144**	2	0.011
3	**0.199**	3	0.169	**3**	**0.055**
				4	0.199
Lags	3	Lags	3	Lags	3
Lags AIC	3	Lags AIC	3	Lags AIC	3
Lags HQ	2	Lags HQ	2	Lags HQ	1
Lags SIC	1	Lags SIC	2	Lags SIC	1
Hungary					
H0 : r	p-value	H0 : r	p-value	H0 : r	p-value
0	0	0	0	0	0
1	0	1	0.000	1	0.005
2	0.023	2	0.014	**2**	**0.060**
3	**0.183**	**3**	**0.624**	3	0.266
4	0.411	4	0.818	4	0.549
				5	0.850
Lags	**1**	Lags	1	**Lags**	**3**
Lags AIC	3	Lags AIC	3	Lags AIC	2
Lags HQ	1	Lags HQ	1	Lags HQ	1
Lags SIC	1	Lags SIC	1	Lags SIC	1

Appendix D Baseline VECM- Error correction terms by country

Figure 11: CZE

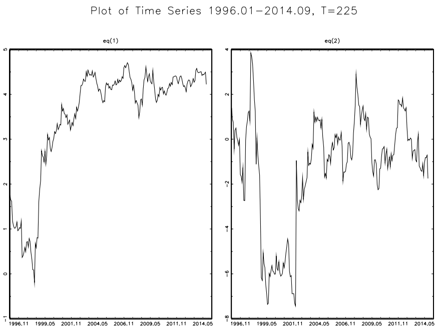


Figure 12: HRV

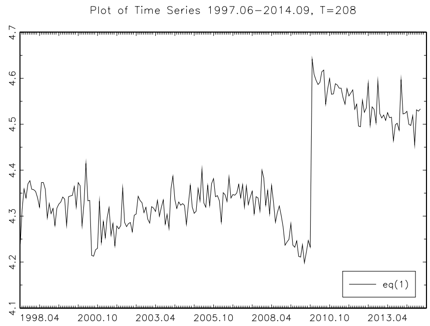


Figure 13: BGR

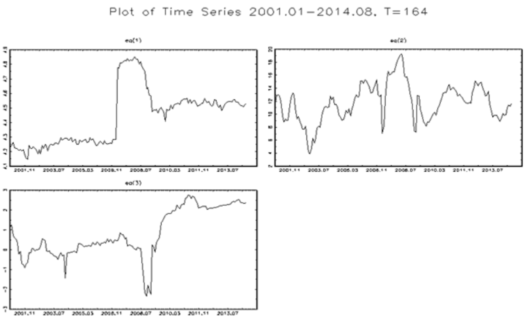


Figure 14: HUN

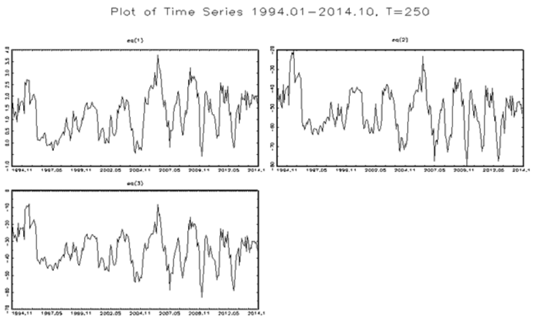


Figure 15: POL

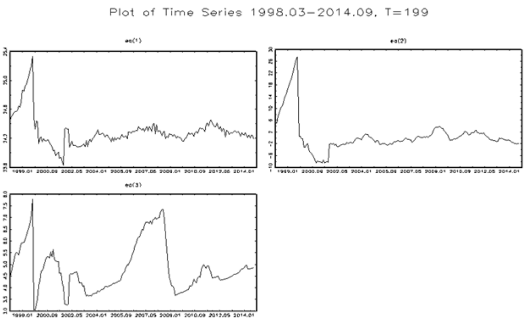
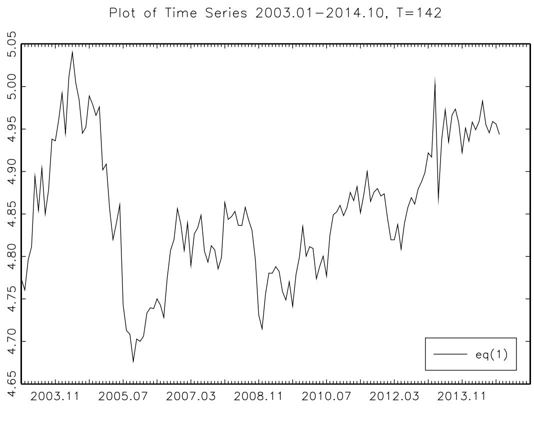


Figure 16: ROU



Appendix E No breaks VECM- Error correction terms by country

Figure 17: CZE

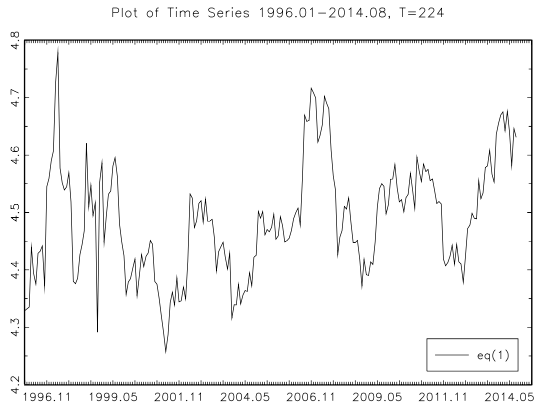


Figure 18: BGR

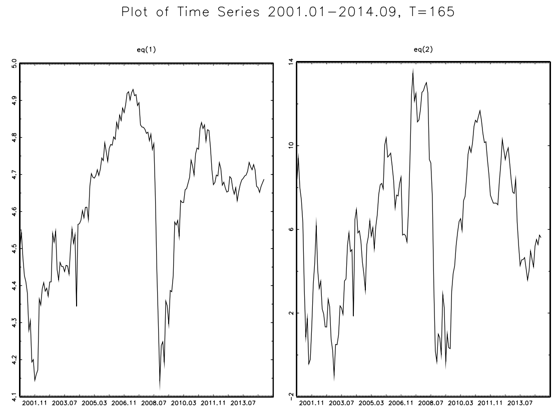


Figure 19: HUN

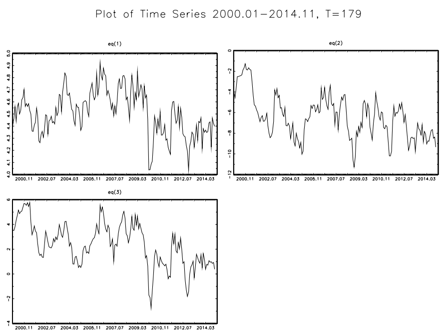


Figure 20: POL

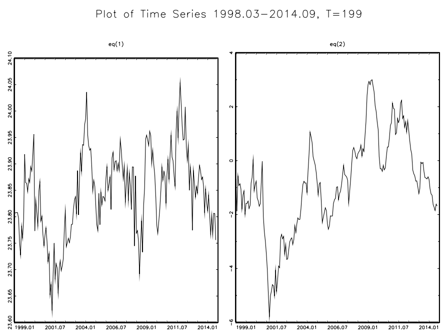
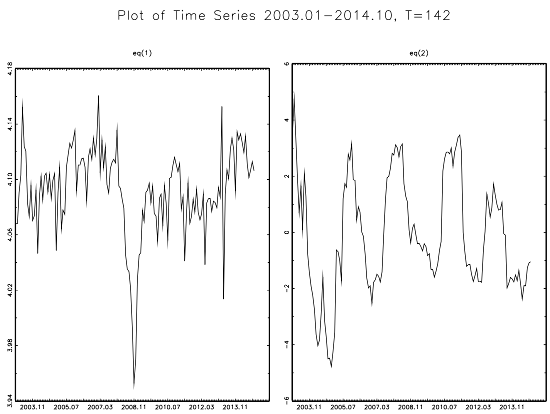


Figure 21: ROU



Appendix F Inflation VECM- Error correction terms by country

Figure 22: CZE

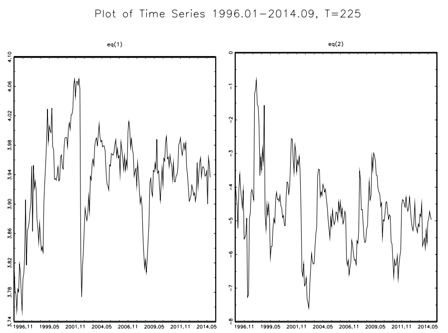


Figure 23: HRV

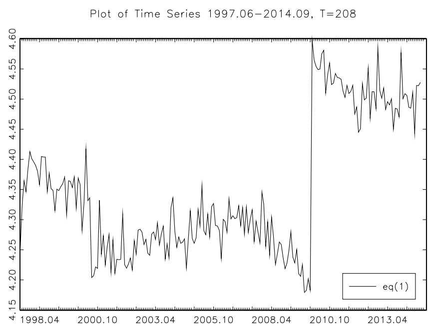


Figure 24: BGR

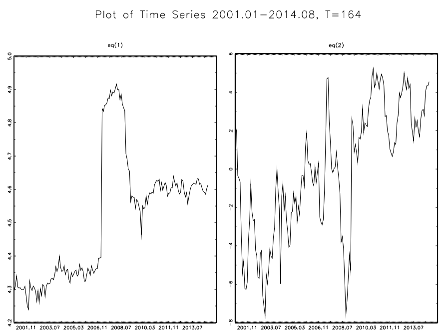


Figure 25: HUN

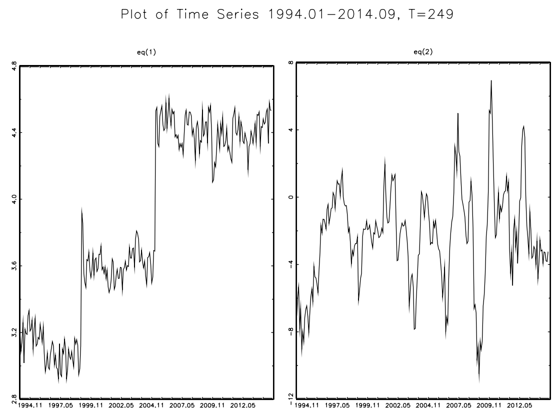
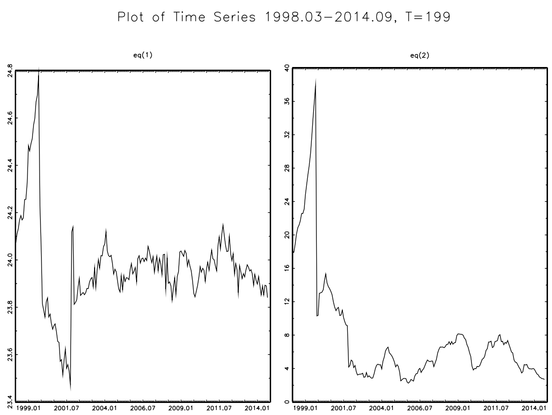
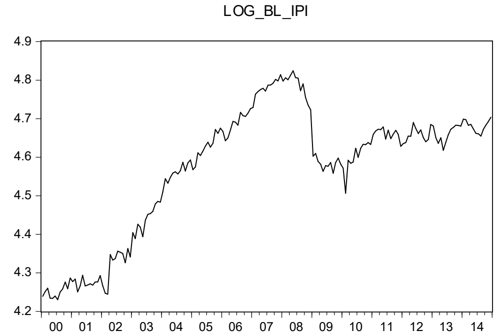
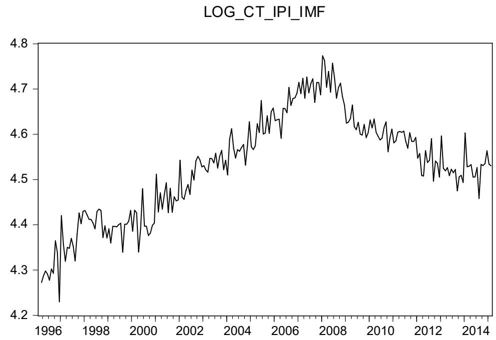


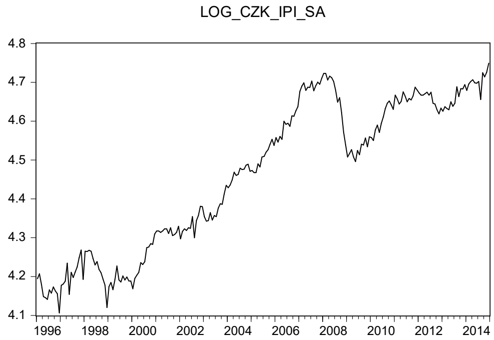
Figure 26: POL

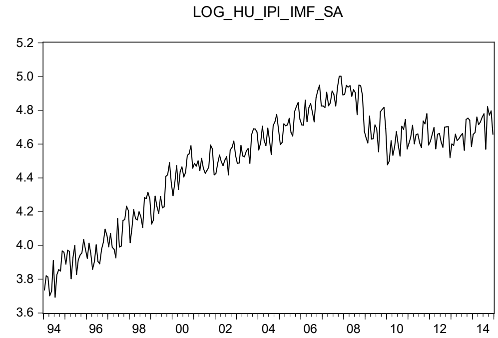


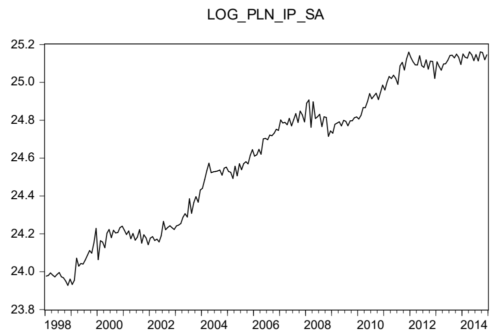
Appendix G Industrial Production series by country after seasonal adjustment and log transformation

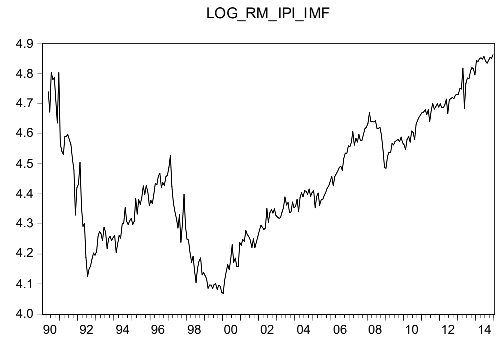












Appendix H Euro area policy rates, Euro Interbank Offered Rate (EONIA) and Euro area Shadow Short Rate (Krippner, 2015)

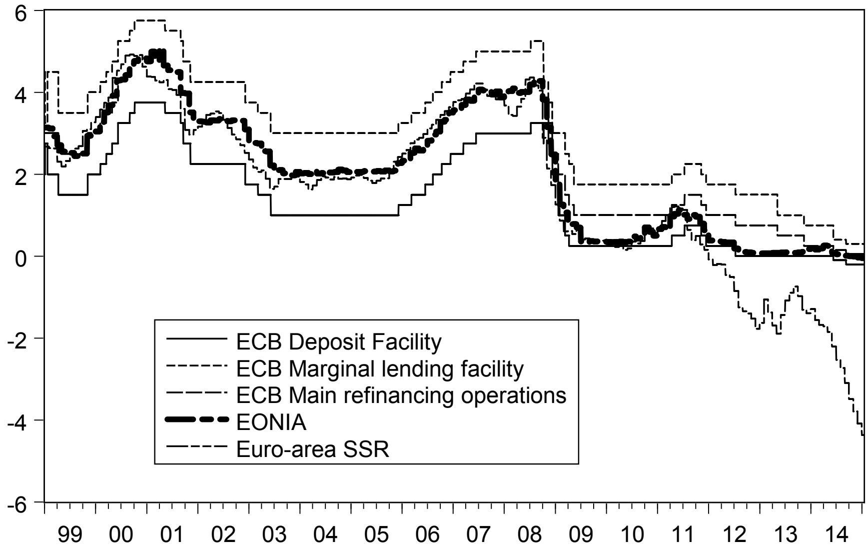


Table 10: Euro area policy, SSR and interbank rates - Correlation matrix - 1999:M1 - 2014:M12

	ECB DEPOSIT FACILITY	ECB MARGINAL LENDING	ECB_MRO	EONIA	EURO AREA SSR
ECB_DEPOSIT_FACILITY	1.000	0.940	0.978	0.977	0.860
ECB_MARGINAL_LENDING	0.940	1.000	0.988	0.951	0.935
ECB_MRO	0.978	0.988	1.000	0.973	0.920
EONIA	0.977	0.951	0.973	1.000	0.887
EURO_AREA_SSR	0.860	0.935	0.920	0.887	1.000

List of Tables